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Financial Market Integration and the Mobility of Capital: Evidence from the OECD Countries

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Abstract

This study examines the long-run relationship between domestic saving and investment and takes a country-by-country account of the mobility of capital and integration of international financial markets. The analysis is carried out for a comprehensive set of 24 OECD countries. The study finds support for the cointegrating relationship between domestic saving and investment for a number of countries. The slope parameter of saving remains well above zero for most countries. The support for cointegration between domestic saving and investment suggests the sustainability of current account deficits and the solvency of intertemporal budget constraint. The degree of capital mobility and the integration of financial markets vary across countries. The reliance on domestic saving in the countries with low to moderate mobility of capital underlines the need to accelerate domestic saving to finance the accumulation of capital and keep the current account imbalances in sustainable bounds. The investment in the countries with high mobility of capital is financed by a world pool of capital. The major concerns for the countries with high mobility of capital are the vulnerability to the speculative (systematic or stochastic) expectations (rational or irrational) of international investors, sustainability of current account deficits, adequacy of foreign exchange reserves, and the stability of the financial system.

JEL Classification: C30, E21, E22, F21, F32, F41.

Keywords: Saving, Investment, Capital mobility, Cointegration, Structural breaks

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1. Introduction

The globalization of financial markets and the mobility of capital across countries remain an area of unresolved controversy in the open economy macroeconomics. An archival assessment of the behavior of international financial markets suggests that these markets broke up during World War I, made a brief comeback during 1925-31, and then withered in the Great Depression. The macroeconomic policy choices tend to be constrained by 'Policy Trilemma' in that a country can, at the most, choose two (any two) of the three policy objectives: independent monetary policy, fixed exchange rate system, and open capital account. The 'Policy Trilemma' or 'Impossible Trinity' during the Bretton Woods system required the imposition of policy restrictions on capital flows to support the fixed exchange rates and enable the central banks to conduct autonomous monetary policies. The major economies of the world remained integrated only by the most rudimentary, and typically bilateral, trade and financial arrangements until 1950. The private capital movements began to return in the 1960s, grew rapidly in the 1970s and then grew even faster since the 1980s. The very success of the Bretton Woods system in spurring international trade and related capital movements brought about its own collapse by resurrecting the 'Inconsistent Trinity' (Obstfeld 1998). The Mundell-Fleming model (Mundell 1962, 1963, Fleming 1962) postulates the perfect mobility of capital and frictionless integration of international financial markets. Any differential between the domestic and foreign interest rates is offset by the inflow (outflow) of foreign capital into (out of) the domestic economy, and is reflected in (i) the accumulation (depletion) of foreign exchange reserves under the fixed exchange rate regime and (ii) an appreciation (depreciation) of domestic currency under the flexible exchange rate regime. The switch from 'fixed' to 'floating' exchange rate system since the early-1970s led to uncertainty in the foreign exchange markets and spurred financial risks in investing in foreign assets. The policy-driven devaluation and market-driven depreciation of the external value of foreign currency reduces the rates of return on foreign financial assets and induces home-bias in the asset portfolios of domestic investors.

The influential paradigm pioneered by Feldstein and Horioka (1980) postulates the absence of perfect mobility of capital, presence of home-bias in the asset portfolios of investors, and near-autarkic behaviour of international financial markets. Feldstein and Horioka (FH) estimate the reduced-form cross-sectional regression of domestic investment on saving (both scaled by GDP) for the OECD countries (1960-74), and find the slope parameter of saving to be significantly different from zero, but not from unity. They interpret these results in terms of the long-run immobility of capital and imperfect integration of international capital markets. Several studies have reinforced the FH conclusions and systematically shown the presence of high correlation between the rates of saving and investment or equivalently the lack of correlation between current account and rate of investment (Feldstein 1983, Murphy 1984, Penati and Dooley 1984, Feldstein and Bacchetta 1991, Tesar 1991, Coakley et al. 1996, Jansen and Schulze 1996, Coiteux and Olivier 2000, De Vita and Abbott 2002, Caporale et al. 2005, Singh 2008, 2013). The empirical evidence nevertheless remains mixed and inconclusive in that a number of studies have contrarily provided weak or no support for the FH hypothesis and high SI correlations (Yamori 1995, Obstfeld and Rogoff 1996, Blanchard and Giavazzi 2002, Kejriwal 2008, Byrne et al. 2009, Giannone and Lenza 2009, Kumar and Rao 2011, Ketenci 2012, 2013). Several factors could be catalytic to the imperfect mobility of capital and home-bias in the asset portfolios of investors, such as the informational inefficiencies in the international financial markets, bounded rationality of investors, exchange rate risks, low risk-adjusted rates of return on foreign assets, moral hazards in debt and equity markets, barriers to trade, and the transportation and transaction costs.

The FH interpretation of high SI correlations in terms of the low mobility of capital came to be viewed at variance with the observed high mobility of capital as manifested by large capital flows, competitive returns on financial assets in international capital markets, and persistent current account imbalances in the OECD countries. The micro-founded intertemporal optimization approach to current account, which came into vogue contemporaneously with the FH strand since the early-1980s, generally accepts the findings of numerically high and statistically significant SI correlations in FH strand, but develops several theoretical channels to explain these correlations in the wake of high international mobility of capital (IMC). It interprets high SI correlations as the corollary of current account solvency constraint, rather than as an index of capital immobility (Glick and Rogoff 1995, Obstfeld and Rogoff 1996, Obstfeld 1998). The high SI correlations in the wake of high IMC could arise from several factors such as the productivity and technology shocks, global shocks, common factors, current account solvency constraints, endogenous government policy responses to current account imbalances, economic growth, fiscal deficits, and large country size; see Singh (2007) for a survey. The intertemporal budget constraint may not allow countries to run high and perpetual current account deficits and the solvency constraint requires the long-run relationship between domestic saving and investment. The SI correlations, as such, tend to be high regardless of the degree of capital mobility and integration of financial markets across countries.

The financial globalization facilitates the diversification of investment portfolios across international financial assets, global sharing of financial risks, maximization of risk-adjusted rates of returns, and the efficient allocation of world capital resources. The advantage of exchange rate certainty that was lost with the switch from ‘fixed’ to ‘floating’ exchange rate regime since the breakdown of the Bretton Woods system in the early-1970s has partially been compensated by the development of information technology and innovations of financial derivative products. The information technology has helped reduce (though not completely remove) the asymmetric information problem, and that development of risk-management financial derivative products (forwards, futures, options, and swaps) has helped hedge (though not perfectly hedge) the exposures to exchange rate risks. The technological and financial innovations in conjunction with the liberalization of capital account and development of financial sector contributed to the decrease in home-bias in the asset portfolios of investors and the increase in mobility of capital across countries. Obstfeld (1998) argues that the worldwide trends in financial opening in the 1990s have restored a degree of IMC not seen since the beginning of this century. The integration of international financial markets has increased due to the financial innovations and liberalization, technological breakthrough, and the growth of world trade. The Maastricht Treaty (February 1992) and the formation of the European Union provided an added dimension and possibly contributed to the integration of financial markets among the member countries. The gains of financial openness nevertheless remain surrounded by the risks of financial liberalization. The recurrent episodes of financial crises suggest that the sudden stops and panic reversals of high-resolution and speculative capital inflows could make the economies prone to financial calamities with self-fulfilling runs on currencies. The crises in a given economy could cascade across countries and destabilize even the informationally efficient and financially solvent systems.

This study examines the long-run relationship between domestic saving and investment, and measures the international mobility of capital and integration of financial markets for 24 OECD countries. An analysis of the cointegrating relationship between domestic saving and investment is essentially inevitable to determine the validity of intertemporal budget constraint and assess the sustainability of current account deficits. The saving-investment behaviour and implied SI correlations tend to differ across countries. The parameter (average) estimates obtained from the cross-sectional and panel data models provide information only for the sample-group as a whole and, as such, tend to lose relevance for the formulation and assessment of country-specific economic policies. The study uses the time-series approach and takes a country-by-country account of the SI correlations and IMC. The long-run model is estimated using several estimators to assess the robustness of results across methodologies and test statistics. Most studies conducted in a time- series econometric setting have assumed a temporally stable behaviour of financial markets and examined SI correlations without allowing structural breaks in the cointegrating vector. The financial markets are vulnerable to the speculative (systematic or stochastic) expectations (rational or irrational) of international investors and, as such, are susceptible to structural breaks and regime switches. The structural breaks reduce the power of cointegration tests and weaken the robustness of statistical evidence obtained from one-regime models with time-invariant parameters and no structural break. The study allows structural breaks in the cointegrating vector and takes a robust account of the SI correlations. The remainder of the study is organised as follows. Section 2 specifies the model. Section 3 presents the empirical results. Section 4 undertakes an analytical account of the interpretations of SI correlations and discusses the policy implications of results. Section 5 sums up the conclusions.

2. The Model

The reduced-form bi-variate FH model is used to estimate the long-run relationship between domestic saving and investment and measure the international mobility of capital.

$$(1) \qquad [I/Y]_t = \alpha + \beta[S/Y]_t + \varepsilon_t; \quad \varepsilon(t)\sim \mathrm{iid}\left(0,\sigma^2\right); \quad t \in [1,\dots,T]$$

The $\partial[I/Y]/\partial[S/Y] = \beta \in [0,1]$ in model (1) is the saving-retention coefficient and it shows the proportion of incremental saving retained and invested in the country of origin. The higher the proportion of saving retained in the domestic economy, the lower the proportion is lent and invested in the international financial markets. In a complete financial autarky with $\beta = 1$, all domestic saving are retained and invested in the country of origin and there are no borrowing and lending across countries and, as such, no international mobility of capital. In contrast when $\beta = 0$, all domestic saving are lent and invested in international capital markets and there is frictionless mobility of capital and, thus, perfect integration of international financial markets. The investment in such case is fully financed by a world pool of capital, rather than domestic saving. The intermediate state with $0 < \beta < 1$ characterises the current account imbalances and moderate mobility of capital.

Model (1) is estimated on annual data with a constant time space $T \in \{1970, 1971, \dots, 2006\}$ for 24 OECD countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Japan, Luxembourg, The Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, Turkey, the U.K., and the U.S. The sample period encompasses major macroeconomic events in the global economy that contributed to the integration of goods and financial markets across countries. The 1970s witnessed the breakdown of the Bretton Woods system of fixed exchange rate and a move towards flexible exchange rate regime. The flexible exchange rate system unleashed the 'Impossible Trinity' constraint and provided higher degree of freedom for the formulation of economic policies. The countries could dismantle restrictions on capital accounts and, thus, allow capital flows concurrently with the independent monetary policy. A number of countries undertook the liberalization (gradual) of capital accounts, financial sector and stock market since the 1970s and the 1980s (see Kaminsky and Schmukler 2003). The Maastricht Treaty (February 1992) and the formation of the European Union provided an added dimension and plausibly contributed to financial integration among the member countries since the 1990s.

The saving, S, in model (1) is measured in terms of the gross saving, investment, I, in terms of the gross capital formation (gross fixed capital formation plus inventories) and output, Y, in terms of the gross domestic product (GDP); all at current prices. The ratios of saving, $[S/Y]$, and investment, $[I/Y]$, each to GDP (Y) are expressed in per cent and, thus, represent the respective rates of saving and investment (hereafter saving and investment or SI). The economic rationale for using gross, rather than net, saving is two-fold. First, it is gross, rather than net, saving that flows across countries. Second, the methods and accounting practices used for computing deprecation (capital consumption), required to arrive at net saving and investment, differ across countries. The net saving and investment are, therefore, not strictly comparable across countries. All the data used in the study are sourced from the World Development Indicators (Online), The World Bank.

3. Empirical Results

3.1 Unit Root Tests

The unit root tests are first performed to examine the time-series properties of the model series. Such an analysis assumes particular importance for the FH model, as, *prima facie*, it may seem puzzling to recognise the I(d) property of the ratio of two possibly I(d) series of saving (and investment) and GDP; where the order of integration $d \geq 1$ (see Levy 2000, Singh 2008). Since the rates of saving and investment are bounded between zero and one, they may be persistent, rather than I(1) processes. It is, however, a common practice to model such bounded persistent series as I(1), rather than a stationary process (Kejriwal 2008). Nicolau (2002) argues that while it is not possible to say that these bounded time-series are random walks, some of these series behave just like random walks. The paths of these bounded random walks are almost indistinguishable from the usual random walks, although these are stochastically bounded by an upper and lower finite limit. Cavaliere (2005) develops an asymptotic theory for the integrated and near-integrated time-series with some constrained range, and shows that the presence of such constraints can lead to drastically different asymptotics. Kejriwal (2008) simulates the critical values corresponding to the bounded unit root distribution for a set of sample countries, and finds that the critical values are the same as those of standard unit root tests. Hence, the 0–1 bounds on the saving and investment shares are not constraining in any way (Kejriwal 2008).

The augmented Dickey-Fuller (ADF) model (Dickey and Fuller 1981) does not reject the null hypothesis of a unit root for the level as well as first-differenced series of saving and investment for most countries (Table 1). The Phillips-Perron (PP) test (Phillips and Perron 1988) predominantly does not reject the null hypothesis of a unit root for the level, but rejects the null hypothesis for the first-differenced series of saving and investment across most countries. The KPSS (Kwiatkowski, Phillips, Schmidt, and Shin) test (Kwiatkowski et al. 1992) rejects the contrary null hypothesis of no unit root for the level series for some countries. The KPSS test does not test reject the null hypothesis for the first-differenced series of saving and investment ubiquitously for all the sample countries. The conventional ADF and PP tests are known to have low power and that KPSS test a tendency to over-reject the null hypothesis in small samples. The asymptotically powerful DF-GLS and DF-GLSu

(Elliott et al. 1996, Elliott 1999) tests based on the generalised least squares (GLS) are carried out to cross-examine the evidence. The DF-GLS and DF-GLSu tests do not reject the null hypothesis of a unit root for the level series of saving and investment for most countries. These tests counter-intuitively do not reject the null hypothesis for even the first-differenced series for several countries.

The one-regime unit root tests become mis-specified and are not very informative of non-stationarity in the presence of structural breaks in the underlying series. These tests are biased towards non-rejection of the null hypothesis of a unit root, if the underlying series contains a structural break (Perron 1989). The stationary series may erroneously appear to be non-stationary due to the false non-rejection of the null hypothesis. Perron (1989) provides a test for the null hypothesis of a unit root in the presence of a exogenously determined structural break in the series at the known location. The estimates from the Perron (1989) test, however, would be biased in favour of the rejection of the null hypothesis, as the break-point is not treated as data-dependent and unknown under the alternative hypothesis (Zivot and Andrews 1992). The study uses the endogenous structural break unit root tests of Zivot and Andrews (1992), Lumsdaine and Papell (1997) and Lee and Strazicich (2003, 2004) to test the null hypothesis of a unit root and determine the break-points endogenously from the data. These tests involve the estimation of the model for different break dates using the recursive (rolling or sequential) approach, and then performing the grid-search to locate the most significant break-point, $\pi \in \Pi \subset (0, 1)$, endogenously from the data. The observations are trimmed symmetrically from both beginning and end of the sample space, and the trimmed interval of $\Pi = \{0.15 \times \mathrm{T}, 0.85 \times \mathrm{T}\}$ is used to perform the grid-search and locate the break-point.

The Zivot-Andrews test tests the joint null hypothesis of a unit root with no structural break against the alterative hypothesis of a one-time break in the series. It sets $\mathrm{y}(\mathrm{t}) = \alpha + \mathrm{y}(\mathrm{t}-1) + \varepsilon(\mathrm{t})$ as the null model and uses the minimum t-test statistic (highest absolute t-test, $|\mathrm{t}|$, statistic) to test the null hypothesis of a unit root with no structural break. The one-time break, however, could be inadequate and lead to the loss of information in the presence of multiple breaks in the series. Ben-David et al. (2003) argue that just as the failure to allow one break can cause non-rejection of the unit root null by the ADF test, the failure to allow for two breaks, if they exist, can cause non-rejection of the unit root null by the tests which only incorporate one break. They further argue that allowing for more breaks does not necessarily mean more rejections of the unit-root hypothesis, because the critical value increases in absolute value with the inclusion of more breaks (Ben-David et al. 2003). Lumsdaine and Papell (1997) extend the Zivot-Andrews test to accommodate up to two structural breaks in the series at the unknown locations. The Lumsdaine-Papell test allows structural breaks only under the alternative hypothesis and not under the null hypothesis analogous to the Zivot-Andrews test. It follows that the rejection of the null hypothesis in both Zivot-Andrews and Lumsdaine-Papell tests does not necessarily imply the rejection of the unit root per se, but would imply the rejection of a unit root without breaks (Lee and Strazicich 2003). The assumption of no break un-

Table 1

Unit Root Tests with No Structural Break

Country	Conventional Tests						GLS–Based Point Optimal Tests			
	ADF		PP		KPSS		DF–GLS		DF–GLSu	
	I/Y	S/Y	I/Y	S/Y	I/Y	S/Y	I/Y	S/Y	I/Y	S/Y
	Level Series						De–trended Level Series			
Australia	–3.02 (1)	–0.05 (8)	–2.88	–1.96	0.108	0.154**	–2.91** (0)	–0.99 (8)	–3.06	–1.09
Austria	–1.91 (2)	–1.94 (0)	–3.03	–1.78	0.153**	0.190**	–1.26 (6)	–1.03 (8)	–1.32	–1.14
Belgium	–1.28 (5)	–2.48 (8)	–3.30	–2.80	0.182**	0.176**	–0.96 (9)	–1.16 (5)	–1.13	–1.24
Canada	–0.33 (8)	–0.63 (9)	–2.64	–1.84	0.114	0.142	–1.46 (9)	–1.28 (9)	–1.40	–1.28
Denmark	–1.56 (8)	–2.41 (9)	–2.00	–2.02	0.160**	0.122	–1.85 (9)	–1.22 (9)	–4.58*	–1.21
Finland	–1.87 (8)	–1.20 (9)	–2.62	–2.73	0.076	0.104	–2.69 (0)	–2.30 (5)	–2.69	–2.30
France	–0.79 (9)	–3.73** (9)	–1.97	–1.73	0.155**	0.178**	–1.26 (9)	–1.09 (9)	–1.33	–1.15
Germany	–1.31 (8)	–4.83* (9)	–2.81	–2.98	0.102	0.150**	–1.52 (9)	–1.84 (9)	–1.61	–1.94
Greece	–2.64 (9)	–2.35 (9)	–2.79	–2.83	0.110	0.108	–0.93 (8)	–0.86 (8)	–0.94	–0.86
Iceland	1.30 (9)	–2.40 (9)	–0.50	–3.98**	0.195**	0.118	–1.19 (9)	–2.93** (9)	–1.32	–2.72
Ireland	–1.61 (9)	–1.63 (9)	–1.38	–2.16	0.152**	0.173**	–1.99 (9)	–1.19 (9)	–1.95	–1.17
Italy	–0.17 (9)	–2.15 (9)	–3.95**	–3.37	0.125	0.087	–1.93 (9)	–1.73 (9)	–1.83	–1.68
Japan	–2.40 (9)	–3.30 (9)	–2.64	–2.32	0.089	0.100	–1.99 (9)	–2.74 (9)	–2.02	–2.77
Luxembourg	–0.90 (9)	–2.02 (9)	–2.89	–3.24	0.095	0.128	–1.63 (9)	–2.02 (9)	–1.53	–2.07
Netherlands	–1.87 (9)	–1.80 (9)	–3.24	–2.03	0.109	0.121	–1.56 (9)	–2.07 (9)	–1.61	–2.07
New Zealand	–0.46 (9)	–1.84 (9)	–3.11	–3.53	0.094	0.053	–0.88 (9)	–2.08 (8)	–0.87	–2.13
Norway	0.17 (9)	–1.03 (9)	–2.70	–1.40	0.070	0.151**	–1.00 (9)	–1.40 (9)	–1.00	–1.36
Portugal	–2.72 (9)	–1.90 (9)	–2.64	–2.73	0.062	0.074	–2.45 (9)	–1.60 (8)	–2.45	–1.58
Spain	–1.18 (9)	–1.81 (9)	–1.19	–1.56	0.166**	0.177**	–1.00 (9)	–0.72 (9)	–1.05	–0.75
Sweden	–2.13 (9)	–1.97 (9)	–2.91	–2.13	0.068	0.135	–2.07 (9)	–0.90 (9)	–2.04	–0.96
Switzerland	–2.11 (9)	–2.71 (9)	–2.42	–1.94	0.072	0.109	–1.81 (9)	–1.76 (9)	–1.92	–1.90
Turkey	–1.92 (9)	–0.97 (9)	–2.68	–2.06	0.131	0.137	–1.44 (9)	–1.34 (9)	–1.47	–1.37
United Kingdom	–2.63 (9)	–2.07 (9)	–2.88	–3.06	0.053	0.058	–1.93 (9)	–1.94 (9)	–1.83	–1.90
United States	–1.19 (9)	–2.24 (9)	–2.98	–2.92	0.090	0.062	–1.53 (9)	–2.16 (9)	–1.51	–2.16

	First-Differenced Series						First-Differenced De-trended Series			
Australia	–5.67* (1)	–5.62* (0)	–7.35*	–6.04*	0.081	0.062	–6.56* (0)	–5.77* (0)	–6.56*	–5.73*
Austria	–7.34* (0)	–6.36* (0)	–8.65*	–7.19*	0.077	0.068	–7.56* (0)	–6.54* (0)	–7.56*	–6.53*
Belgium	–8.48* (0)	–1.17 (9)	–10.72*	–10.20*	0.063	0.115	–8.69* (0)	–1.24 (9)	–8.70*	–1.25
Canada	–6.45* (0)	–1.64 (9)	–6.88*	–5.00*	0.100	0.080	–6.51* (0)	–1.36 (9)	–6.57*	–1.33
Denmark	–6.01* (0)	–2.20 (9)	–6.36*	–4.07**	0.048	0.092	–15.07* (0)	–2.20 (9)	–14.25*	–2.19
Finland	–4.67* (0)	–2.69 (5)	–4.88*	–5.58*	0.073	0.060	–4.81* (0)	–2.75 (5)	–4.81*	–2.77
France	–2.03 (9)	–1.66 (9)	–6.83*	–4.51*	0.055	0.068	–2.08 (9)	–1.64 (9)	–1.96	–1.68
Germany	–1.10 (9)	–1.90 (9)	–4.18**	–4.88*	0.100	0.132	–1.00 (9)	–1.47 (9)	–0.94	–1.47
Greece	–3.64** (1)	–3.63** (1)	–7.64*	–7.70*	0.046	0.045	–3.73* (1)	–3.71* (1)	–3.74*	–3.73*
Iceland	–2.70 (9)	–2.24 (9)	–6.38*	–9.05*	0.136	0.086	–0.87 (9)	–2.36 (8)	–0.73	–2.08
Ireland	–2.14 (9)	–1.25 (9)	–5.61*	–6.03*	0.107	0.116	–1.75 (9)	–2.20 (8)	–1.75	–2.15
Italy	–2.27 (8)	–1.89 (8)	–10.05*	–7.13*	0.079	0.065	–3.24** (9)	–0.87 (9)	–3.16	–0.95
Japan	–1.67 (8)	–2.08 (9)	–4.71*	–4.80*	0.088	0.115	–1.12 (9)	–1.91 (9)	–1.14	–1.98
Luxembourg	–2.27 (9)	–2.71 (9)	–8.01*	–8.40*	0.093	0.125	–1.91 (9)	–1.57 (8)	–1.64	–1.74
Netherlands	–1.85 (8)	–2.38 (9)	–5.06*	–5.18*	0.107	0.082	–0.71 (9)	–2.35 (9)	–0.70	–2.22
New Zealand	–1.75 (9)	–2.52 (9)	–6.77*	–5.69*	0.062	0.067	–1.59 (9)	–2.15 (9)	–1.46	–2.13
Norway	–4.65* (9)	–1.76 (8)	–5.12*	–4.97*	0.058	0.107	–1.82 (9)	–1.94 (6)	–2.02	–1.86
Portugal	–2.57 (9)	–1.80 (9)	–3.80**	–4.47*	0.054	0.060	–2.44 (9)	–1.50 (9)	–2.35	–1.50
Spain	–2.56 (9)	–3.15 (9)	–4.02**	–4.72*	0.055	0.057	–2.58 (9)	–3.149** (9)	–2.65	–3.11
Sweden	–2.32 (8)	–3.00 (9)	–5.22*	–4.42*	0.069	0.063	–1.62 (9)	–2.45 (8)	–1.61	–2.41
Switzerland	–1.85 (8)	–10.42* (9)	–4.25*	–4.14**	0.065	0.073	–1.14 (9)	–1.59 (9)	–1.07	–1.41
Turkey	–1.93 (9)	–2.18 (8)	–7.81*	–5.60*	0.068	0.089	–1.81 (9)	–1.57 (8)	–1.77	–1.70
United Kingdom	–2.11 (9)	–2.48 (9)	–5.32*	–6.03*	0.055	0.049	–2.22 (9)	–2.23 (9)	–2.17	–2.02
United States	–1.97 (9)	–2.14 (9)	–6.02*	–5.88*	0.095	0.072	–1.32 (9)	–1.55 (9)	–1.41	–1.65

Notes: (1) All the tests are performed including a constant and a trend in the model; (2) The figures in round parentheses are the autoregressive (AR) lags; (3) The maximal lag is set at $k_{max} = \text{int}\left\{12(T/100)^{1/4}\right\} = 9$ (Schwert 1989), and the modified Akaike information criterion (MAIC) is used to truncate the AR lag-length for the (i) ADF tests performed on the given series of the rates of saving and investment and the (ii) GLS-based point optimal DF-GLS and DF-GLSu tests carried out on the detrended series of the rates of saving and investment; (4) The DF-GLSu test is carried out using the same number of AR lags of the detrended series as used for the DF-GLS test; (5) The results obtained from the PP and KPSS tests performed on the model estimated with one to four lags generally provided similar evidence for the null hypothesis. The results are, therefore, reported only for lag four for the PP and KPSS tests; (6) * and ** indicate the statistical significance and implied rejection of the null hypothesis at 1 % and 5 % levels, respectively. All the tests test the null hypothesis of a unit root, except for the KPSS test which tests the contrary null hypothesis of no unit root in the series.

der the null hypothesis often leads to 'spurious rejections' of the null hypothesis. Lee and Strazicich (2003, 2004) develop the minimum LM unit root tests and allow up to two structural breaks under the null hypothesis[1]. The break-point is determined where the LM t-statistic, obtained from all the possible regressions, is at its minimum (maximum in absolute, |t|, term). The study performs the Lee-Strazicich test for both one and two structural breaks, using the 'break model'. The results suggest that the structural break unit root tests do not reject the null hypothesis of a unit root for the level series, but reject the null hypothesis for the first-differenced series, for most countries (Table 2 and Table 3). These tests generally point towards the I(1) properties of the model series. The break dates (years) are approximately consistent across tests for most countries.

3.2 Tests for Cointegration and the Long-Run Estimates

3.2.1 Standard OLSEG Estimates

The OLS-based two-step estimator of Engle and Granger (OLSEG) (1987) is first used to examine the long-run relationship between the model series. The OLSEG sequentially involves the estimation of the static regression model in levels, $y(t) = \alpha + \beta x(t) + \varepsilon(t)$, and then the estimation of an auxiliary, $\varepsilon(t) = \gamma\varepsilon(t-1) + \nu(t)$, or augmented auxiliary, $\Delta\varepsilon(t) = \gamma\varepsilon(t-1) + \sum_{i=1}^{k} \zeta(i)\Delta\varepsilon(t-i) + \nu(t)$, regression to perform unit root tests on a common stochastic process $\varepsilon(t) = y(t) - \alpha - \beta x(t)$ and test $H_0{:}\varepsilon(t) \sim I(1)$ (no cointegration among I(1) variables) against $H_1{:}\varepsilon(t) \sim I(0)$ (cointegration among I(1) variables). The OLSEG estimates reject the null hypothesis of no cointegration between saving and investment for Australia, Japan, Portugal, Spain, Turkey and the U.K., but not for the remaining set of countries. The estimated slope parameter of saving is dimensionally (i) small ($0 \leq \beta < 0.50$) and implied IMC is high for Denmark, Iceland, Luxembourg, Ireland, The Netherlands, Norway, Sweden, the U.K. and the U.S., (ii) moderate ($0.50 \leq \beta < 0.75$) and implied IMC is moderate for Canada, New Zealand, Portugal and Turkey, and (iii) large ($0.75 \leq \beta \leq 1$) and implied IMC is low for Australia, Austria, Belgium, Finland, France, Germany, Greece, Italy, Japan, Spain and Switzerland (Table 4).

[1] Lee and Strazicich (2003, 2004) allow for (i) an abrupt change in level, but no change in trend rate, in the 'crash model' (denoted as Model A) and (ii) simultaneous changes in both level and trend in the 'break model' (denoted as Model C). The asymptotic null distribution of the two-break LM unit root test (i) is invariant to the location ($\lambda = T_B/T$) and magnitude of structural break in the model with one-time change in level (Model A) and (ii) depends on the location ($\lambda = T_B/T$) of structural break in the model with a change in both level and trend (Model C). The critical values in Model C (model with intercept and trend break) are symmetric around λ and $(1-\lambda)$. The study performs the Lee-Strazicich tests, for both one and two structural breaks, using the 'break model' (Model C).

Table 2

Unit Root Tests with One Structural Break

Country	Zivot-Andrews		Lumsdaine-Papell		Lee-Strazicich	
	I/Y	S/Y	I/Y	S/Y	I/Y	S/Y
	Level Series					
Australia	–5.22** (1) [1991]	–3.27 (1) [1992]	–5.25** (1) [1990]	–3.21 (1) [1994]	–3.81 (1) [1999]	–2.41 (8) [1988]
Austria	–5.09** (5) [1982]	–4.83** (0) [1975]	–4.95 (5) [1981]	–3.98 (0) [1979]	–5.43* (7) [1988]	–4.10 (8) [1987]
Belgium	–6.24* (0) [1988]	–6.05* (0) [1987]	–6.33* (0) [1987]	–6.21* (0) [1986]	–6.21* (6) [1983]	–3.95 (8) [1985]
Canada	–4.01 (0) [1991]	–5.47* (1) [1990]	–3.98 (0) [1990]	–5.44** (1) [1989]	–3.87 (0) [1990]	–3.37 (8) [1990]
Denmark	–4.15 (3) [1980]	–4.71 (1) [1983]	–4.12 (0) [1979]	–5.10** (1) [1982]	–5.57* (6) [1987]	–3.74 (1) [1983]
Finland	–3.95 (9) [2001]	–4.95** (1) [1997]	–4.18 (9) [2000]	–4.93 (1) [1996]	–3.65 (9) [1991]	–4.47** (1) [1995]
France	–1.57 (7) [1998]	–5.66* (3) [1981]	–1.66 (7) [1998]	–5.30** (3) [1980]	–3.69 (6) [2002]	–6.997* (4) [1986]
Germany	–5.30** (3) [1990]	–3.31 (8) [1986]	–5.27** (3) [1989]	–3.67 (8) [1985]	–3.13 (1) [1990]	–4.19 (8) [1984]
Greece	–3.58 (9) [1997]	–3.58 (9) [1993]	–3.76 (9) [1996]	–3.55 (9) [1996]	–5.65* (7) [1986]	–4.34 (9) [1986]
Iceland	–1.41 (6) [1994]	–5.20** (0) [1982]	–1.64 (6) [1992]	–5.24** (0) [1981]	–2.95 (7) [1984]	–4.04 (0) [1984]
Ireland	–3.41 (0) [1983]	–3.63 (1) [1979]	–3.40 (0) [1982]	–3.66 (1) [1978]	–4.00 (9) [1990]	–4.59** (7) [1993]
Italy	–4.54 (5) [1996]	–4.63 (2) [1995]	–4.28 (5) [1995]	–4.62 (2) [1994]	–4.76** (4) [1994]	–6.89* (7) [1993]
Japan	–5.55* (1) [1988]	–4.57 (9) [1988]	–5.59* (1) [1987]	–4.50 (9) [1987]	–3.15 (8) [1986]	–4.44 (9) [1986]
Luxembourg	–4.02 (0) [1978]	–4.10 (6) [1989]	–3.96 (0) [1977]	–4.13 (6) [1994]	–5.23* (8) [1985]	–3.70 (6) [1998]
Netherlands	–3.22 (6) [1988]	–4.06 (0) [1975]	–3.18 (6) [1987]	–3.70 (0) [1980]	–4.32 (6) [1986]	–3.21 (5) [1987]
New Zealand	–3.74 (0) [1987]	–5.12** (5) [1982]	–3.74 (0) [1989]	–5.21** (5) [1981]	–5.16* (5) [1988]	–4.26 (2) [1989]
Norway	–4.92 (2) [1997]	–3.06 (2) [1999]	–4.91 (2) [1996]	–2.95 (2) [1998]	–2.94 (1) [1987]	–5.16* (1) [1989]
Portugal	–5.56* (1) [1983]	–6.04* (1) [1978]	–5.56** (1) [1982]	–6.40* (1) [1977]	–5.38* (4) [1983]	–3.62 (6) [1983]
Spain	–3.18 (9) [1997]	–4.46 (2) [1979]	–3.23 (9) [1997]	–4.59 (2) [1978]	–4.34 (6) [1992]	–5.50* (7) [1992]
Sweden	–4.20 (1) [1993]	–5.50* (5) [1992]	–4.20 (1) [1992]	–5.54** (5) [2001]	–3.18 (5) [1985]	–4.21 (5) [1984]
Switzerland	–4.59 (1) [1979]	–3.98 (1) [1984]	–4.80 (1) [1978]	–3.98 (1) [1983]	–3.41 (1) [1983]	–2.59 (8) [1987]
Turkey	–4.80** (0) [1987]	–7.59* (2) [1987]	–4.76 (0) [1986]	–7.38* (2) [1986]	–3.91 (7) [1996]	–5.15* (1) [1986]
United Kingdom	–6.71* (6) [1999]	–4.35 (2) [1997]	–6.68* (6) [1998]	–4.36 (2) [1996]	–5.36* (6) [1998]	–3.92 (1) [1985]
United States	–3.49 (0) [1990]	–4.06 (5) [1996]	–3.47 (0) [1987]	–4.04 (5) [1995]	–3.89 (8) [1989]	–3.38 (3) [1986]

Continued next page

Table 2 (continued)

Country	Zivot-Andrews I/Y	Zivot-Andrews S/Y	Lumsdaine-Papell I/Y	Lumsdaine-Papell S/Y	Lee-Strazicich I/Y	Lee-Strazicich S/Y
	First Differenced Series					
Australia	–6.54* (1) [1991]	–6.41* (7) [1992]	–6.64* (1) [1990]	–5.46** (7) [1991]	–6.53* (8) [1986]	–6.47* (6) [1988]
Austria	–7.53* (1) [1985]	–6.66* (0) [1984]	–7.46* (1) [1984]	–7.47* (0) [1976]	–5.86* (1) [1984]	–5.58* (0) [1984]
Belgium	–8.82* (0) [1987]	–9.40* (0) [1984]	–8.68* (0) [1986]	–9.25* (0) [1983]	–8.04* (8) [1994]	–5.44* (8) [1994]
Canada	–7.04* (0) [1975]	–6.43* (1) [1994]	–7.13* (0) [1983]	–6.36* (1) [1993]	–5.99* (0) [1983]	–5.16* (2) [1992]
Denmark	–5.60* (8) [1984]	–6.81* (1) [1982]	–5.90* (8) [1991]	–6.61* (1) [1981]	–4.90** (8) [1985]	–4.75** (8) [1997]
Finland	–4.91** (5) [1990]	–6.17* (1) [1994]	–4.87 (5) [1989]	–6.34* (1) [1993]	–4.30 (0) [1989]	–4.42 (1) [1996]
France	–4.63 (6) [1992]	–4.92** (0) [1983]	–5.02 (6) [1994]	–5.11** (0) [1982]	–5.01** (6) [1995]	–4.73** (4) [1986]
Germany	–5.66* (3) [2001]	–5.89* (0) [1976]	–5.61* (3) [1987]	–6.26* (0) [1975]	–5.53* (3) [1999]	–5.86* (4) [1990]
Greece	–3.60 (8) [1992]	–3.92 (8) [1992]	–7.89* (8) [1986]	–9.00* (8) [1986]	–4.89** (7) [1994]	–5.64* (7) [1994]
Iceland	–4.71 (5) [1981]	–8.84* (1) [1979]	–5.18** (5) [1984]	–8.69* (1) [1978]	–5.06** (0) [1986]	–7.61* (1) [1985]
Ireland	–6.69* (0) [1980]	–6.57* (0) [1982]	–6.60* (0) [1979]	–6.60* (0) [1981]	–4.28 (5) [1985]	–5.60* (3) [1996]
Italy	–7.35* (1) [1977]	–4.49 (0) [1980]	–7.68* (1) [1978]	–4.86 (2) [1982]	–7.67* (0) [1988]	–4.90** (3) [1989]
Japan	–6.52* (1) [1992]	–6.07* (0) [1992]	–6.46* (1) [1991]	–5.90* (0) [1991]	–4.71** (5) [1990]	–4.571** (1) [1992]
Luxembourg	–9.41* (0) [1976]	–4.16 (9) [1994]	–9.89* (0) [1975]	–2.51 (9) [2000]	–5.76* (0) [1995]	–4.569** (9) [1985]
Netherlands	–4.86** (5) [1983]	–5.55* (2) [1980]	–4.63 (5) [1982]	–5.24** (2) [1979]	–5.17* (7) [1988]	–5.62* (0) [2002]
New Zealand	–7.51* (0) [1975]	–4.38 (8) [1989]	–6.78* (0) [1978]	–5.14** (8) [1988]	–5.39* (8) [1990]	–5.12* (2) [1984]
Norway	–6.04* (9) [1991]	–7.11* (1) [1985]	–5.77* (9) [1990]	–6.995* (1) [1984]	–5.10** (7) [1995]	–6.41* (1) [1987]
Portugal	–4.91** (5) [1983]	–4.93** (0) [1976]	–4.83 (5) [1982]	–8.42* (0) [1975]	–6.29* (5) [1995]	–4.49 (7) [1988]
Spain	–4.59 (8) [1995]	–5.38* (7) [1984]	–4.44 (8) [1994]	–6.28* (7) [1991]	–4.06 (8) [1990]	–4.98** (2) [1989]
Sweden	–5.49* (0) [1990]	–5.33** (1) [1990]	–5.40** (0) [1989]	–5.23** (1) [1989]	–4.83** (6) [1985]	–5.09** (1) [1996]
Switzerland	–5.18** (1) [1978]	–5.05** (0) [1991]	–5.69* (1) [1976]	–5.74* (0) [1976]	–5.82* (1) [1999]	–6.00* (2) [1989]
Turkey	–7.54* (0) [1989]	–6.43* (1) [1989]	–7.95* (0) [1997]	–7.20* (1) [1988]	–7.31* (0) [1998]	–5.04** (1) [1999]
United Kingdom	–5.44* (6) [1990]	–6.13* (0) [1976]	–6.24* (6) [1989]	–6.98* (0) [1978]	–5.16* (1) [1991]	–5.05** (0) [1993]
United States	–5.91* (0) [1992]	–5.82* (0) [1980]	–5.80* (0) [1991]	–5.85* (0) [1998]	–5.18* (0) [1989]	–5.46* (0) [1997]

Notes: (1) All the tests are performed including a constant and a trend in the model; (2) The figures in square brackets are the break-points (years); (3) The figures in round parentheses are the AR lags. The maximal lag is set at $k_{max} = \text{int}\left\{12(T/100)^{1/4}\right\} = 9$ (Schwert 1989), and the AR lag-length is determined using the (i) Akaike information criterion (AIC) in the Zivot-Andrews and Lumsdaine-Papell tests and the (ii) general-to-specific approach in the Lee-Strazicich test. The lags are sequentially dropped from the end in the general-to-specific approach, and the significance level of 10% is used as the minimum cut-off to trim the lags; (4) * and ** indicate the statistical significance at 1% and 5% levels, respectively.

Table 3

Unit Root Tests with Two Structural Breaks

Country	Lumsdaine-Papell		Lee-Strazicich	
	I/Y	S/Y	I/Y	S/Y
	Level Series			
Australia	–6.09 (1) [1978, 1990]	–4.86 (1) [1982, 1990]	–5.13 (5) [1986, 1997]	–5.12 (1) [1986, 1990]
Austria	–8.66* (5) [1981, 2001]	–4.83 (0) [1978, 1984]	–7.31* (1) [1986, 2002]	–6.25** (9) [1987, 1992]
Belgium	–7.23* (0) [1981, 1987]	–7.36* (0) [1984, 1990]	–8.62* (5) [1985, 1993]	–9.16* (6) [1983, 1994]
Canada	–6.66 (0) [1981, 1990]	–6.06 (1) [1986, 1994]	–4.14 (2) [1988, 1996]	–7.10* (8) [1990, 2001]
Denmark	–5.57 (3) [1980, 1990]	–6.63 (1) [1980, 1992]	–8.61* (6) [1987, 1996]	–6.75* (7) [1993, 1999]
Finland	–5.93 (9) [1987, 2000]	–6.77** (1) [1990, 2001]	–9.76* (5) [1987, 1991]	–6.30** (7) [1990, 1999]
France	–4.23 (7) [1982, 2000]	–5.46 (3) [1987, 1998]	–4.99 (6) [1985, 1995]	–7.71* (4) [1986, 1993]
Germany	–7.73* (3) [1987, 2000]	–5.09 (8) [1986, 1992]	–7.84* (8) [1986, 1999]	–5.72** (8) [1984, 1999]
Greece	–7.25* (9) [1988, 1996]	–6.33 (9) [1988, 1993]	–9.09* (4) [1983, 1992]	–9.42* (4) [1983, 1992]
Iceland	–3.67 (6) [1990, 1996]	–6.44 (0) [1975, 1986]	–3.90 (2) [1987, 1999]	–5.09 (8) [1985, 2000]
Ireland	–4.20 (0) [1980, 1988]	–4.72 (1) [1978, 1996]	–5.67 (7) [1984, 1998]	–6.91* (7) [1989, 1995]
Italy	–4.96 (5) [1989, 1995]	–6.71 (2) [1980, 1994]	–8.73* (9) [1984, 1992]	–6.69* (2) [1989, 1996]
Japan	–6.70 (1) [1978, 1987]	–3.996 (9) [1984, 1989]	–8.70* (8) [1986, 1999]	–5.21 (8) [1986, 1999]
Luxembourg	–6.73 (0) [1976, 1992]	–4.22 (6) [1993, 1998]	–7.56* (8) [1985, 1992]	–4.13 (6) [1998, 2002]
Netherlands	–6.03 (6) [1983, 1999]	–5.04 (0) [1980, 1987]	–6.48* (7) [1986, 2000]	–4.89 (8) [1987, 2000]
New Zealand	–4.95 (0) [1975, 1986]	–10.16* (5) [1981, 1989]	–8.13* (5) [1989, 1997]	–7.39* (8) [1988, 2000]
Norway	–5.87 (0) [1985, 1996]	–4.87 (2) [1979, 1995]	–6.33** (7) [1990, 2000]	–7.13* (1) [1984, 1990]
Portugal	–6.42 (1) [1976, 1982]	–8.80* (1) [1977, 1990]	–6.51* (4) [1983, 1995]	–6.05** (7) [1984, 1994]
Spain	–3.49 (9) [1993, 1999]	–6.84** (2) [1978, 1991]	–6.53* (6) [1986, 1992]	–7.26* (7) [1984, 2000]
Sweden	–6.36 (1) [1979, 1992]	–6.19 (5) [1980, 1990]	–6.85* (6) [1986, 1991]	–6.67* (6) [1993, 1997]
Switzerland	–6.89** (1) [1978, 1991]	–5.78 (1) [1978, 1991]	–7.00* (6) [1985, 1997]	–4.78 (0) [1983, 1990]
Turkey	–5.65 (0) [1986, 2000]	–7.53* (2) [1979, 1986]	–6.87* (8) [1986, 1994]	–10.25* (6) [1985, 1994]
United Kingdom	–8.11* (6) [1987, 1998]	–5.13 (2) [1980, 1996]	–10.26* (6) [1986, 1994]	–5.76** (9) [1989, 1995]
United States	–4.61 (0) [1976, 1993]	–4.06 (5) [1985, 2001]	–5.18 (5) [1984, 1996]	–7.32* (7) [1984, 1997]

Continued next page

Table 3 (continued)

Country	Lumsdaine-Papell I/Y	Lumsdaine-Papell S/Y	Lee-Strazicich I/Y	Lee-Strazicich S/Y
	First Differenced Series			
Australia	–7.45* (1) [1990, 1996]	–7.47* (7) [1983, 1999]	–6.11** (1) [1989, 1993]	–8.27* (6) [1984, 1997]
Austria	–9.01* (1) [1976, 1984]	–7.77* (0) [1976, 1983]	–6.08** (1) [1984, 1999]	–6.72* (9) [1988, 1995]
Belgium	–9.34* (0) [1980, 1989]	–9.55* (0) [1980, 1989]	–11.09* (6) [1983, 1993]	–14.52* (7) [1983, 1992]
Canada	–8.69* (0) [1981, 1989]	–6.91** (1) [1977, 1993]	–7.00* (0) [1983, 1992]	–6.37** (2) [1983, 1992]
Denmark	–6.78** (8) [1988, 1996]	–7.60* (1) [1981, 1998]	–7.53* (8) [1985, 1995]	–7.35* (8) [1983, 1997]
Finland	–5.42 (5) [1990, 2001]	–7.47* (1) [1990, 1996]	–6.60* (0) [1989, 1994]	–5.60 (9) [1984, 1994]
France	–5.95 (6) [1982, 1994]	–5.61 (0) [1979, 1989]	–8.19* (9) [1994, 1999]	–8.15* (8) [1986, 1998]
Germany	–6.78** (3) [1980, 1987]	–6.76* (0) [1975, 1982]	–9.70* (9) [1984, 1999]	–7.57* (4) [1986, 1999]
Greece	–8.84* (8) [1986, 1995]	–9.98* (8) [1986, 1995]	–11.18* (4) [1983, 1992]	–11.43* (4) [1983, 1992]
Iceland	–7.02** (5) [1983, 1989]	–9.24* (1) [1978, 1992]	–7.36* (0) [1985, 2001]	–7.86* (1) [1986, 2003]
Ireland	–6.95** (0) [1979, 2000]	–7.71* (0) [1981, 1994]	–5.56 (9) [1984, 1995]	–6.995* (7) [1989, 1993]
Italy	–8.04* (1) [1978, 1994]	–5.83 (2) [1979, 1997]	–8.92* (0) [1987, 1994]	–5.59 (2) [1998, 2003]
Japan	–7.187** (1) [1977, 1991]	–6.747 (0) [1976, 1991]	–5.79** (5) [1990, 1997]	–8.57* (9) [1990, 1999]
Luxembourg	–10.41* (0) [1975, 1996]	–4.96 (9) [1987, 1999]	–7.58* (5) [1984, 1992]	–7.27* (8) [1984, 2000]
Netherlands	–6.00 (5) [1987, 1996]	–5.49 (2) [1978, 1985]	–7.29* (9) [1984, 1999]	–6.01** (1) [1986, 1991]
New Zealand	–7.24* (0) [1980, 1991]	–6.58 (8) [1985, 1998]	–8.65* (8) [1988, 2000]	–5.90** (2) [1991, 2003]
Norway	–10.52* (9) [1985, 1997]	–7.39* (1) [1979, 1988]	–6.28** (9) [1983, 2001]	–7.07* (1) [1985, 2002]
Portugal	–6.53 (5) [1982, 1996]	–9.48* (0) [1975, 1993]	–8.15* (8) [1993, 2003]	–6.09** (0) [1985, 1994]
Spain	–4.77 (8) [1991, 1997]	–7.87* (7) [1991, 2000]	–6.53* (6) [1988, 1997]	–6.16** (2) [1991, 1998]
Sweden	–6.07 (0) [1975, 1989]	–6.90** (1) [1982, 1993]	–7.99* (3) [1988, 1993]	–6.41** (7) [1984, 1994]
Switzerland	–6.35 (1) [1976, 1990]	–7.16** (0) [1976, 1990]	–6.48* (6) [1989, 1998]	–7.03* (2) [1983, 1989]
Turkey	–8.69* (0) [1977, 1988]	–9.34* (1) [1982, 1988]	–8.58* (0) [1986, 2002]	–7.21* (1) [1986, 1991]
United Kingdom	–7.23* (6) [1983, 1989]	–7.91* (0) [1975, 1981]	–6.37** (1) [1992, 1998]	–6.43** (3) [1985, 1998]
United States	–6.77** (0) [1975, 1983]	–6.74 (0) [1981, 1998]	–6.32** (0) [1992, 2001]	–6.08** (0) [1993, 2001]

Notes: (1) All the tests are performed including a constant and a trend in the model; (2) The figures in square brackets are the break-points (years); (3) The figures in round parentheses are the AR lags. The maximal lag is set at $k_{max} = \text{int}\left\{12(T/100)^{1/4}\right\} = 9$ (Schwert 1989), and the AR lag-length is determined using the (i) AIC in the Lumsdaine-Papell test and the (ii) general-to-specific approach in the Lee-Strazicich test. The lags are sequentially dropped from the end in the general-to-specific approach, and the significance level of 10% is used as the minimum cut-off to trim the lags; (4) * and ** indicate the statistical significance at 1% and 5% levels, respectively.

Table 4

OLSEG Estimates of the Long-Run Model

		t–Ratios				
		With Corrected Standard Errors@		Without Corrected Standard Errors		ADF Statistics
Country	β	$H_0{:}\beta=0$	$H_0{:}\beta=1$	$H_0{:}\beta=0$	$H_0{:}\beta=1$	$H_0{:}\varepsilon(t)\sim I(0)$
Australia	0.84	10.45*	–2.01**	11.56*	–2.23**	–4.16*
Austria	0.81	8.31*	–1.91**	6.14*	–1.41	–0.40
Belgium	0.89	12.80*	–1.58	9.96*	–1.23	–1.32
Canada	0.64	6.47*	–3.67*	4.59*	–2.60*	–1.85
Denmark	0.05	0.33	–5.86*	0.30	–5.32*	–2.15
Finland	1.28	5.01*	1.08	4.75*	1.03	–1.60
France	0.95	14.17*	–0.75	10.45*	–0.56	–1.88
Germany	0.84	4.97*	–0.95	4.44*	–0.84	–0.48
Greece	0.99	55.70*	–0.69	52.61*	–0.65	–1.55
Iceland	0.72	3.82*	–1.46	3.87*	–1.48	–1.51
Ireland	–0.05	–0.80	–17.83*	–0.65	–14.54*	–1.57
Italy	1.13	5.72*	0.65	5.54*	0.63	–2.62
Japan	1.01	30.35*	0.30	23.32*	0.23	–4.04*
Luxembourg	0.24	2.90*	–9.05*	3.13*	–9.76*	–2.44
Netherlands	0.49	2.49**	–2.54*	2.49*	–2.54*	–2.70
New Zealand	0.56	2.16**	–1.70**	2.06**	–1.62	–2.97
Norway	–0.36	–1.52	–5.73*	–1.24	–4.67*	–1.84
Portugal	0.58	5.03*	–3.61*	3.39*	–2.44*	–3.50**
Spain	1.19	7.995*	1.27	6.74*	1.07	–3.35***
Sweden	0.45	2.05**	–2.54*	1.92**	–2.38**	–1.40
Switzerland	1.53	18.73*	6.48*	11.45*	3.96*	–0.97
Turkey	0.71	8.90*	–3.65*	7.60*	–3.12*	–3.42**
United Kingdom	0.49	6.94*	–7.19*	4.25*	–4.40*	–3.22***
United States	0.38	4.32*	–6.98*	4.26*	–6.88*	–2.67

Notes: (1) All the models are estimated including the intercept (constant) term. The estimates of the intercept term are not reported to conserve space; (2) @ The heteroscedasticity-robust estimates of the standard errors are obtained using the Eicker-White heteroscedasticity-consistent estimator. (3) *, ** and *** indicate the statistical significance and implied rejection of the null hypothesis at 1%, 5% and 10% levels, respectively. (4) The ADF statistics are for the null hypothesis of a unit root in the residuals, $\varepsilon(t)$, of model (1). The critical values for the null hypothesis of a unit root in $\varepsilon(t)$ and implied null hypothesis of no cointegration between the model series are (i) 3.96 at 1% and 3.37 at 5% level (Hamilton 1994) and (ii) 3.39 at 1% level and 2.76 at 5% level (Phillips and Ouliaris 1990).

3.2.2 GMM Estimates

The standard OLSEG estimates become biased and inefficient in the presence of non-orthogonality of regressors and serial-correlation of residuals. The 'super-consistency' property of OLS indeed allows one to omit I(0) regressors from the cointegrating model and asymptotically ignore the problems of endogeneity of regressors and serial-correlation of residuals[2]. In small samples, however, the OLS estimates remain biased and have inferential problems for the significance of long-run parameters. The bias is often substantial (Banerjee et al. 1993, Inder 1993) and the t statistics of the cointegrating coefficients are generally not valid for statistical inference. The study uses the generalised method of moments (GMM) estimator of Hansen (1982) to resolve the problem of endogeneity of regressors and obtain the efficient parameter estimates. The GMM has desirable properties in large samples. Consider a linear regression model, $y = X\beta + \mu$, with the underlying moment conditions given by $E\mu\mu'] = I\sigma^2$, $E[X'\mu] = E[X'(y–X\beta)] = 0$ and $V[X'\mu] = [X'X]\sigma^2$; where E is the expectation operator, X the matrix of exogenous variables and μ the vector of residuals. The OLS estimators are obtained by minimizing $\mu'\mu = (y - X\beta)'(y - X\beta)$. The OLS parameter estimates should satisfy the moment condition in terms of the orthogonality (zero correlation) between exogenous variables and error term, such that $E[X'\mu] = 0$. This moment condition, $E[X'\mu] = 0$, could either come from the assumptions about the model variables or/and be derived from the first-order conditions of an optimization problem. This condition is operationalized by replacing the expectation operator, E, by the sample average, $\frac{1}{T}\sum_{t=1}^{T} X'_t(y_t - X_t\beta) = 0$. The moment conditions in the method of moments (MM) are equal to the number of unknown parameters.

If the orthogonality condition is not satisfied such that $E[X'\mu] \neq 0$, then it is essential to replace X by some instrument set, Z, and minimize $E[Z'\mu]$ so that $E[Z'\mu] = 0$ and $V[Z'\mu] = [Z'Z]\sigma^2$; where Z is assumed to have the same dimension as X. The instrumental variables (IV) estimators, $\hat{\beta}_{IV} = (Z'X)^{-1}Z'y$, are obtained by setting $Z'\mu = 0$ or $Z'(y - X\beta) = 0$. If there are more moment conditions than the number of parameters, then the system of equations becomes algebraically over-identified and cannot be solved. This situation arises because the lagged (twice-lagged, thrice-lagged, ...) variables tend to be the weak instruments, and it often becomes necessary to have a large set of moment conditions. Sargan (1958) develops the generalised instrumental variables estimator and suggests minimizing $(y - X\beta)'Z(Z'Z)^{-1}Z'(y - X\beta) = \mu'Z(Z'Z)^{-1}Z'\mu$. The matrix $(Z'Z)^{-1}$ serves to weight the orthogonality conditions. The weighting matrix comes into operation when the dimensions of Z are larger than the dimensions of X. Hansen (1982) shows that a more efficient estimator can be obtained by replacing $(Z'Z)^{-1}$ by the optimal weighting in terms of the inverse of the matrix, $MCOV(z\mu) = \sum_{k=-L}^{L}\sum_{t=1}^{T} Z'_t\mu_t\mu_{t-k}Z_{t-k}$. The study performs GMM estimation using the optimal weights suggested by Hansen (1982). The $J \sim \chi^2(n - p)$ statistic is used to test the moment conditions and examine the non-orthogonally of regressors to the residual process; where n is the number of instruments and p the num-

[2] The OLS estimators of a regression are 'super-consistent' when the model series are cointegrated. Instead of approaching their true values at a rate proportional to $n^{-1/2}$, the OLS estimates will approach them at a rate proportional to n^{-1} (Davidson and Mackinnon 1993).

ber of parameters. The GMM reduces to MM estimator when the number of parameters are equal to the number of moment conditions.

The GMM estimates of the model suggest that the t-ratios reject: (i) $H_0{:}\beta = 0$, but not $H_0{:}\beta = 1$, for Australia, Belgium, Canada, Finland, Germany, Greece, Iceland, Italy, Japan and Portugal, (ii) $H_0{:}\beta = 1$, but not $H_0{:}\beta = 0$, for Denmark, Ireland, Luxembourg, The Netherlands and Norway, (iii) both $H_0{:}\beta = 0$ and $H_0{:}\beta = 1$ for Switzerland, Turkey, the U.K. and the U.S. and (iv) neither $H_0{:}\beta = 0$ nor $H_0{:}\beta = 1$ for Austria, France, New Zealand, Spain and Sweden (Table 5). The *J*-statistics do not reject the null hypothesis of no correlation between instruments

Table 5

GMM Estimates of the Long-Run Model

Country	β	t-Ratios		*J*-Statistics
		H_0: β=0	H_0: β=1	
Australia	0.82	6.30**	–1.39	2.31 (0.13)
Austria	0.88	5.38	–0.72	1.07 (0.30)
Belgium	1.02	8.70*	0.16	1.53 (0.22)
Canada	0.61	3.60*	–2.26	1.67 (0.20)
Denmark	0.06	0.25	–4.06*	0.15 (0.70)
Finland	1.45	2.73**	0.85	1.63 (0.20)
France	1.03	10.44	0.26	1.01 (0.31)
Germany	0.67	2.14**	–1.04	2.14 (0.14)
Greece	0.97	33.59*	–1.15	1.40 (0.24)
Iceland	0.78	2.24**	–0.62	1.30 (0.25)
Ireland	–0.02	–0.24	–11.01**	0.55 (0.46)
Italy	1.19	4.25*	0.69	2.22 (0.14)
Japan	1.06	20.49*	1.10	0.66 (0.42)
Luxembourg	0.15	1.59	–9.32*	0.07 (0.78)
Netherlands	0.35	1.45	–2.66**	0.57 (0.45)
New Zealand	0.81	2.15	–0.52	1.23 (0.27)
Norway	–0.33	–0.90	–3.60*	0.67 (0.41)
Portugal	0.73	5.34**	–1.98	0.24 (0.62)
Spain	1.33	5.99	1.48	4.60 (0.03)
Sweden	0.36	1.21	–2.15	0.56 (0.45)
Switzerland	1.60	11.50*	4.33*	0.39 (0.53)
Turkey	0.64	7.01*	–3.90*	1.33 (0.25)
United Kingdom	0.57	4.37*	–3.30*	0.52 (0.47)
United States	0.40	3.26*	–4.83*	0.66 (0.41)

Notes: (1) All the models are estimated including the intercept (constant) term. The estimates of the intercept term are not reported to conserve space; (2) The instruments used in GMM estimation are a constant and two lags of [S/Y]. The figures in parentheses corresponding to *J*-Statistics are the p-values; (3) * and ** indicate the statistical significance and implied rejection of the null hypothesis at 1% and 5% levels, respectively.

(used for endogenous regressor) and residual term for all the countries. The slope parameter is dimensionally (i) small $(0 \leq \beta < 0.50)$ and implied IMC is high for Denmark, Ireland, Luxembourg, The Netherlands, Norway, Sweden and the U.S. (ii) moderate $(0.50 \leq \beta < 0.75)$ and implied IMC is moderate for Canada, Germany, Portugal, Turkey and the U.K. and (iii) large $(0.75 \leq \beta \leq 1)$ and implied IMC is low for Australia, Austria, Belgium, Finland, France, Greece, Iceland, Italy, Japan, New Zealand, Spain and Switzerland.

3.2.3 Optimal DOLS, NLLS and FMOLS Estimates

The efficiency of the GMM estimation hinges heavily on the quality (weak or strong) and validity (orthogonality) of instruments. The instruments that are weakly related to endogenous regressors (weak instruments) and are non-orthogonal to the Gaussian disturbances (invalid instruments) can still produce biased and inconsistent estimates. The weak instruments may yield biased two-stage least squares (2SLS) estimates even in large samples (Bound et al. 1995, Staiger and Stock 1997). When several regressors in a model are instrumented, then the validity requirements for the instruments used for endogenous regressors become even more stringent (Staiger and Stock 1997). It is, in fact, difficult to find appropriate instruments that are strongly correlated with endogenous regressors, but are uncorrelated with the Gaussian disturbances. An optimal alternative to using standard OLS and GMM estimators is to introduce an explicit AR(1) specification for X(t) along with the stochastic model for the relationship between Y(t) and X(t). The triangular representation of the cointegrated system of Phillips (1991), with I(1) series of Y(t) and X(t) and I(0) series of μ(t), suggests that the OLS estimator of β is consistent, but not generally fully efficient,

$$\text{(2)} \qquad \mathrm{Y(t)} = \alpha + \beta \mathrm{X(t)} + \mu\mathrm{(t)}$$

$$\text{(3)} \qquad \Delta \mathrm{X(t)} = \eta\mathrm{(t)}$$

The asymptotic distribution of OLS estimator depends on various nuisance parameters engendered by serial-correlation in μ(t) and by correlation between μ(t) and the innovation term for ΔX(t) in model (3). The μ(t) and η(t) are cross-correlated not only contemporaneously, but also at various leads and lags. Phillips (1991) suggests using the following representation for μ(t),

$$\text{(4)} \qquad \mu\mathrm{(t)} = \sum\nolimits_{\mathrm{j}=-\mathrm{k}}^{\mathrm{k}} \delta\mathrm{(j)}\eta\mathrm{(t-j)} + \xi\mathrm{(t)}$$

The $\xi(\mathrm{t})$ in model (4) is not correlated with $\eta(\mathrm{t}-\mathrm{j})$, $\forall \mathrm{j} \in [-\mathrm{k}, \mathrm{k}]$. The cointegrating model (2) can be augmented with the lags and leads of ΔX(t) to resolve the problem of cross-correlation between μ(t) and η(t) (Phillips and Loretan 1991, Saikkonen 1991, Stock and Watson 1993). By substituting model (3) into model (4) and then substituting the resulting transform into model (2), the lags and leads cointegration estimator can be represented as

$$\text{(5)} \qquad Y(t) = \alpha + \beta X(t) + \sum\nolimits_{j=-k}^{k} \delta(j) \Delta X(t-j) + \xi(t)$$

Since $\xi(t)$ is not correlated with $\eta(t)$ in model (4), it will also be uncorrelated with $\Delta X(t)$ in model (5). The $\Delta X(t)$ asymptotically eliminates the effect of endogeneity of X(t) on the distribution of OLS estimator of β. If $\xi(t)$ is independently and identically distributed, then the standard distribution theory can be used to perform inference on the OLS parameter estimates. The Y(t) and X(t) have a common stochastic trend if they are cointegrated and, therefore, the dynamic OLS (DOLS) estimator of Saikkonen (1991) and Stock and Watson (1993) is consistent even if X(t) is endogenous. While the lags and leads of $\Delta X(t)$ resolve the problem of endogeneity of X(t), they do not necessarily eliminate all the serial-correlation and heteroscedasticity in $\xi(t)$. Stock and Watson (1989, 1993) suggest using the generalised least squares (GLS) to estimate model (5). The GLS estimates of standard errors and variance-covariance matrix could be used to construct the asymptotically valid χ^2 hypothesis tests on β. Phillips and Loretan (1991) suggest using a parametric correction in model (5) to account for the potential serial-correlation in $\xi(t)$. The requisite information set for valid conditioning is better modelled by using lagged equilibria than by using lagged differences of the dependent variable, and they recommend augmenting model (5) with the lagged levels of $[Y(t) - \alpha - \beta X(t)]$.

$$\text{(6)} \qquad \begin{aligned} Y(t) = \beta_0 + \beta X(t) + \sum\nolimits_{j=-k}^{k} \delta(j) \Delta X(t-j) \\ + \sum\nolimits_{j=1}^{k} \varphi(j)[Y(t-j) - \alpha - \beta X(t-j)] + \zeta(t) \end{aligned}$$

The $\zeta(t)$ is serially uncorrelated and model (6) can be estimated using the nonlinear least squares (NLLS) estimator. The NLLS estimator of β is asymptotically efficient and the estimates of variance-covariance matrix have the standard limiting distribution. The NLLS variance-covariance matrix can be used to perform hypothesis test on β in a standard manner. Both DOLS and NLLS estimators are unbiased and asymptotically efficient in the presence of endogeniety of regressors and serial-correlation of residuals.

Phillips and Hansen (1990) develop the fully-modified OLS (FMOLS) estimator to resolve the problems of endogeneity-bias and serial-correlation, and obtain the efficient estimates of the model parameters. The FMOLS estimator starts with the standard OLS regression and then, analogous to the Phillips-Perron (Phillips and Perron 1988) unit root test, makes a non-parametric correction to account for the endogeneity-bias and serial-correlation that may show up in the OLS residuals. The estimates of the long-run parameters and associated t-statistics are, thus, adjusted to correct for the bias arising from the endogeneity of regressors and serial-correlation of residuals. The FMOLS estimator is super-consistent and is asymptotically both unbiased and normally distributed (Park and Phillips 1988, Phillips and Hansen 1990, Hansen and Phillips 1990). Phillips (1995) shows that the FMOLS estimator is reliable in the case of full rank or cointegrated I(1) regressors as well as I(0) regressors. The t-statistics of the long-run coefficients are asymptotically normally distributed, and the standard limiting distributions can be used to perform statistical inference in the FMOLS estimates.

The model with DOLS and NLLS are estimated using the lags-leads structure (k) of k = {–2, 0, +2} each for the first-differenced dynamic regressors[3]. The model with FMOLS is estimated using the lag-length of four for the Bartlett kernel. The results are generally consistent in terms of both numerical magnitude and statistical significance of the slope parameter of saving across DOLS and FMOLS estimators (Table 6). The DOLS estimates suggest that the t-ratios reject (i) $H_0{:}\beta = 0$, but not $H_0{:}\beta = 1$, for Australia, Canada, Finland, France, Greece, Iceland, Japan, New Zealand, Switzerland, (ii) $H_0{:}\beta = 1$, but not $H_0{:}\beta = 0$, for Denmark, Germany, Ireland, The Netherlands and Sweden, (iii) both $H_0{:}\beta = 0$ and $H_0{:}\beta = 1$ for Austria, Belgium, Italy, Luxembourg, Spain, Turkey, the U.K. and the U.S. and (iv) neither $H_0{:}\beta = 0$ nor $H_0{:}\beta = 1$ for Norway and Portugal (Table 6). The slope parameter of saving is dimensionally (i) small $(0 \leq \beta < 0.50)$ and implied IMC is high for Belgium, Luxembourg, The Netherlands, Norway and the U.S., (ii) moderate $(0.50 \leq \beta < 0.75)$ and implied IMC is moderate for Austria, Canada, Turkey and the U.K. and (iii) large $(0.75 \leq \beta \leq 1)$ and implied IMC is low for Australia, France, Greece and Portugal. Some of the slope coefficients are counter-intuitively either unduly large (more than unity) or unduly small (negative). By and large, similar pattern is suggested by the NLLS and FMOLS estimates (Table 6).

3.2.4 Maximum-Likelihood System Estimates

The maximum-likelihood (ML) system estimator of Johansen (1991) estimates the k^{th} order vector autoregression (VAR) model and takes a system-based account of endogeneity.

$$\Delta X(t) = \sum_{i=1}^{k-1} \Gamma(i)\Delta X(t-i) + \tilde{\Pi} X(t-1) + \mu + \varepsilon(t) \tag{7}$$

3 The use of a uniform lag (and lead) structure for all the countries involves an element of arbitrariness. It, however, needs to be recognised that the use of an over-parameterized model with a larger lag (and lead) structure tends to impinge upon the efficiency and finite sample properties of the least squares estimates, particularly if the sample space is not sufficiently large relative to the number of variables in the model. In contrast, the choice of an over-parsimonious model, with lower lag (and lead) structure, could lead to the problem of serial-correlation in the model residuals. The choice of the model structure with k = {–2, 0, +2} seems reasonable, given the considerations of over-parameterization and parsimony.

Table 6

Optimal DOLS, NLLS and FMOLS Estimates of the Long-Run Model

	DOLS [k = {–2, 0, +2}]			NLLS [k = {–2, 0, +2}]			FMOLS [Bartlett kernel lags = 4]		
		t–Ratios			t–Ratios			t–Ratios	
Country	β	H_0: β=0	H_0: β=1	β	H_0: β=0	H_0: β=1	β	H_0: β=0	H_0: β=1
Australia	0.94	6.70*	–0.46	0.94	7.67*	–0.46	0.84	8.13*	–1.50
Austria	0.50	2.13**	–2.14**	0.75	1.32	–0.44	0.71	2.68*	–1.09
Belgium	0.48	2.90*	–3.12*	–1.57	–0.71	–1.17	0.74	3.71*	–1.33
Canada	0.70	2.26**	–0.99	1.17	1.65	0.23	0.68	2.65*	–1.24
Denmark	–0.18	–0.74	–4.99*	0.23	0.27	–0.91	–0.07	–0.22	–3.53*
Finland	1.29	1.80**	0.40	3.34	1.17	0.82	1.43	2.64*	0.79
France	0.90	3.75*	–0.43	1.05	3.59*	0.18	0.91	5.27*	–0.50
Germany	–0.29	–0.65	–2.93*	–10.39	–0.07	–0.08	0.47	1.13	–1.27
Greece	0.92	17.98*	–1.46	0.91	12.19*	–1.27	0.96	27.62*	–1.04
Iceland	1.15	5.57*	0.73	0.71	1.12	–0.45	0.88	3.05*	–0.43
Ireland	–0.05	–0.38	–7.74*	0.07	0.21	–2.86*	–0.01	–0.10	–7.39*
Italy	1.73	5.26*	2.21**	1.83	3.42*	1.55	1.38	3.93*	1.07
Japan	1.05	15.70*	0.75	1.08	18.48*	1.37	0.98	14.39*	–0.33
Luxembourg	0.25	4.25*	–12.76*	0.20	2.26**	–9.00*	0.21	1.77**	–6.69*
Netherlands	0.09	0.37	–3.76*	0.09	0.14	–1.46	0.35	1.21	–2.23**
New Zealand	2.52	2.42**	1.46	4.05	1.51	1.13	0.66	1.57	–0.81
Norway	0.05	0.06	–1.17	1.33	0.44	0.11	–0.17	–0.31	–2.08**
Portugal	0.75	1.30	–0.44	0.82	0.74	–0.17	0.61	2.26**	–1.45
Spain	1.60	6.95*	2.59*	1.84	5.86*	2.67*	1.29	5.10*	1.14
Sweden	–0.10	–0.25	–2.73*	4.31	0.31	0.24	0.20	0.46	–1.80**
Switzerland	1.45	2.25**	0.70	0.09	0.05	–0.56	1.52	5.63*	1.92**
Turkey	0.72	5.04*	–1.94**	0.71	3.85*	–1.54	0.73	6.19*	–2.34**
United Kingdom	0.55	2.22**	–1.78**	0.57	1.24	–0.93	0.49	2.67*	–2.80*
United States	0.46	3.67*	–4.25*	0.60	1.31	–0.89	0.35	2.51*	–4.68*

Notes: (1) All the models are estimated including the intercept (constant) term. The estimates of the intercept term are not reported to conserve space; (2) * and ** indicate the statistical significance and implied rejection of the null hypothesis at 1% and 5% levels, respectively.

Table 7

Maximum-Likelihood System Estimates of the Long-Run Model [VAR lag k =2]

Country	β	Eigenvalues		λ-trace Test		λ-trace Test@	
		H_0: r = 0	H_0: r $\leq$ 1	H_0: r = 0	H_0: r $\leq$ 1	H_0: r = 0	H_0: r $\leq$ 1
Australia	0.90	0.38	0.19	24.14**	7.50**	22.37**	6.94**
Austria	–0.63	0.17	0.01	6.80	0.20	6.3	0.17
Belgium	–0.15	0.38	0.08	19.57**	2.82	18.10**	2.38
Canada	2.76	0.20	0.11	11.54	3.87**	10.42	3.58
Denmark	–0.52	0.16	0.06	8.52	2.27	7.63	2.08
Finland	30.91	0.22	0.07	11.364	2.66	10.37	2.50
France	–2.38	0.15	0.12	10.24	4.60**	9.26	3.81
Germany	–1.06	0.32	0.00	13.38	0.10	11.85	0.09
Greece	0.83	0.16	0.06	8.40	2.24	7.82	2.00
Iceland	1.57	0.19	0.06	9.51	2.11	8.75	1.65
Ireland	–0.74	0.13	0.02	5.65	0.84	5.17	0.68
Italy	1.64	0.25	0.04	11.46	1.58	10.40	1.52
Japan	1.02	0.41	0.08	21.50**	2.85	19.17**	2.71
Luxembourg	0.33	0.26	0.12	14.92	4.47**	14.11	4.02**
Netherlands	0.01	0.24	0.09	12.90	3.10	11.55	2.81
New Zealand	3.91	0.46	0.13	26.26**	5.04**	24.42**	5.00**
Norway	16.73	0.15	0.09	9.06	3.41	8.14	3.23
Portugal	–0.79	0.28	0.27	22.35**	10.80**	19.87**	10.56**
Spain	1.76	0.32	0.10	17.38**	3.66	15.41**	3.48
Sweden	–1.23	0.38	0.03	17.58**	0.88	15.81**	0.83
Switzerland	–4.50	0.28	0.03	12.29	0.91	10.90	0.86
Turkey	0.85	0.26	0.13	15.48**	4.82**	14.27	4.56**
United Kingdom	0.43	0.24	0.07	12.04	2.36	10.95	2.26
United States	0.26	0.26	0.00	10.54	0.06	9.44	0.06

Notes: (1) The r denotes the number of cointegrating vectors; (2) @ indicates the λ-trace corrected for small sample; (3) The 95% critical values are 15.408 for H_0: r = 0 and 3.841 for H_0: r $\leq$ 1; (4) ** indicates the statistical significance at 5% level; (5) $ Long-run coefficients of the first cointegrating vector are normalised on I/Y.

The $\Gamma(i) = -[I-\sum_{i=1}^{k-1}\Pi(i)]$, $\tilde{\Pi} = -[I-\sum_{i=1}^{k}\Pi(i)], \Pi = \alpha\beta'$, $X' = [\{I/Y\}\{S/Y\}]$ is a $p \times 1$ vector of p number of I(1) variables, μ is a vector of constants and $\varepsilon(t)$ is a p-dimensional vector of disturbances with zero mean and covariance matrix so that $\varepsilon(t) \sim iid(0,\Sigma)$. The VAR model (7) is estimated using a lag structure (k) of k=2 for all the countries. The asymptotic λ-trace and the λ-trace adjusted for small-sample consistently do not reject the null hypothesis of no cointegration for most countries. The results provide support for the presence of cointegration only for some countries (Table 7). The long-run slope parameter of saving, β, obtained by normalising the first cointegrating vector on I/Y is counter-intuitively either unduly large (more than unity) or unduly small (negative) for several countries. The slope parameters with unduly large magnitudes and/or perverse signs do not carry any meaningful economic interpretation.

3.3 Structural Breaks

The one-regime estimates become weak and inefficient in the presence of structural breaks and regime-switches in the model parameters. The structural breaks in SI correlations could arise from several factors including the (i) changes in policy stance and the implied liberalization (imposition) of capital controls, (ii) speculative attacks and self-fulfilling runs on currencies, (iii) sudden stops and panic reversals of capital inflows, and the (iv) cyclical states (booms and recessions) of the goods and financial markets. The adoption of flexible exchange rate system since the breakdown of the Bretton Woods system in the early-1970s alleviated the 'Impossible Trinity' constraint and provided higher degree of freedom for the formulation of economic policies. The countries could remove or reduce policy restrictions on the capital account. The Maastricht Treaty (February 1992) and the formation of the European Union provided an added dimension and plausibly contributed to the integration of financial markets among the member countries. The increased mobility of capital tends to be accompanied by the likelihoods of financial crises arising from several factors including the irrational speculations and over-reactions of investors. It needs to be recognised that the timings of structural breaks in the long-run relationship among variables may not strictly coincide with the timings of policy changes and economic shocks (systematic or stochastic), given that the adjustment process may not be instantaneous and the economic agents normally take time to recognise and adjust to the shocks and policy changes. The speed and degree of adjustment depend on a number of factors including the magnitudes of economic and policy shocks, rigidities in the goods and financial markets, and the process of expectation (adaptive or rational) formation among the economic agents. This section allows structural breaks in the cointegrating vector and cross-examines the preceding evidence obtained from one-regime estimators with time-invariant parameters and no structural breaks.

3.3.1 Standard Tests for Model Instability

The structural breaks could arise either from discrete changes in the population regression coefficients at a distinct date or from a gradual evolution of the coefficients over a longer period of time. If the coefficients change gradually over time, then these coefficients would be similar in the adjacent time periods. A commonly used method to examine the constancy (stability) of the model parameters is to compute the parameter estimates using a rolling regression window of fixed size through the sample period. The rolling regression estimation is first initialized using the fixed window width of a given number of observations, and then the estimation is rolled through the remaining sample by advancing one observation at a time[4]. If the

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4 An alternative to rolling regression estimation is the sequential or recursive estimation. The recursive estimation uses all the data points up to the window width, and the sample period is updated ahead by a given increment.

parameters are truly constant over the entire sample period, then the estimates should not be too different across the rolling regression windows. The coefficients obtained from the local estimations are plotted against time to assess the possible changes in these coefficients over the sample period. The estimates of rolling regression are notably not very efficient in that the sampling error in the rolling regressions could be large, due to the small sample size used in the local estimations. Nevertheless, these estimates provide useful information on the time-profile of the coefficients of the model.

The study estimates the rolling OLS regressions using the moving windows of both five and ten years, so as to take a robust account of the possible inflections and breaks in the slope parameter of saving, β. The standard error bands around the time-trajectories of β are constructed to account for the parameter uncertainty inherent in the estimation process, and map the width of the prediction intervals. The OLS estimates of rolling regressions remained consistent across both the moving windows and are, therefore, reported only for the five-year moving window to conserve space (Figure 1). The time-plots suggest that the slope parameter, β, tends to display intertemporal declines for most countries, except for few countries where it either showed increases (Greece and the U.S.) or did not display any unambiguously clear pattern (Iceland, Japan, Portugal, Spain and the U.K.). The intertemporal declines in the magnitude of β point towards the increase in the international mobility of capital and integration of financial markets over time.

A further analysis is undertaken and the parameter instability tests of Hansen (1992a,b,c) are performed on the model estimated using OLS. These tests test the null hypothesis of no change against the alternative hypothesis of a one-time change in the model parameters. The L, L_C and σ^2_ε statistics of Hansen (1992a,b,c) generally reject the null hypothesis of stability and point towards the instability of the model (Table 8). The L statistics reject the null hypothesis and suggest the instability of the individual parameters for most countries. The Lc statistics cross-validate the evidence and consistently reject the joint null hypothesis of stability for all the countries, except for Australia. In contrast, the error variance, σ^2_ε, statistics do not reject the null hypothesis of model stability for most countries.

3.3.2 OLSGH Test and a Single Structural Break

The residual-based OLS estimator of Gregory and Hansen (OLSGH) (1996) is one of the most commonly used estimators to detect structural breaks in the cointegrating vector. The OLSGH is the direct extension of OLSEG, and it allows one-time break, via dummy variable, in either intercept or both intercept and slope parameters. The break date is unknown, *a priori*, and is determined endogenously by the model. The first step in OLSGH involves the estimation of the static regression models augmented with intercept dummy to account for the level shift (Model I), intercept dummy and a linear trend to assess the level shift with trend (Model II) and

Table 8

OLS Estimates of the Long-Run model and the Hansen Tests for Model Instability

Country	β	Hansen Statistics		
		L	Lc	σ_{ε}^2
Australia	0.84	0.18 (0.30)	0.67 (0.21)	0.03 (0.97)
Austria	0.81	2.36 (0.00)*	3.24 (0.00)*	0.61 (0.02)**
Belgium	0.89	2.50 (0.00)*	3.21 (0.00)*	0.28 (0.15)
Canada	0.64	1.84 (0.00)*	2.15 (0.00)*	0.17 (0.31)
Denmark	0.05	1.69 (0.00)*	2.40 (0.00)*	1.29 (0.00)*
Finland	1.28	2.40 (0.00)*	2.76 (0.00)*	0.67 (0.02)**
France	0.95	1.06 (0.00)*	1.33 (0.01)*	0.22 (0.23)
Germany	0.84	2.90 (0.00)*	3.91 (0.00)*	0.85 (0.01)*
Greece	0.99	1.29 (0.00)*	2.56 (0.00)*	0.18 (0.30)
Iceland	0.72	0.48 (0.05)**	1.55 (0.00)*	0.32 (0.12)
Ireland	–0.05	0.61 (0.02)**	2.31 (0.00)*	0.19 (0.28)
Italy	1.13	0.71 (0.01)*	0.93 (0.07)***	0.12 (0.50)
Japan	1.01	0.28 (0.15)	1.03 (0.04)**	0.73 (0.01)*
Luxembourg	0.24	0.29 (0.15)	0.98 (0.05)**	0.34 (0.11)
Netherlands	0.49	2.04 (0.00)*	2.64 (0.00)*	0.40 (0.07)***
New Zealand	0.56	0.90 (0.00)*	1.18 (0.02)**	0.28 (0.15)
Norway	–0.36	2.80 (0.00)*	3.56 (0.00)*	0.46 (0.05)**
Portugal	0.58	033 (0.11)	1.42 (01)*	0.46 (0.05)**
Spain	1.19	0.54 (0.03)**	1.03 (0.04)**	0.15 (0.38)
Sweden	0.45	2.90 (0.00)*	3.58 (0.00)*	0.36 (0.09)***
Switzerland	1.53	1.83 (0.00)*	2.40 (0.00)*	0.19 (0.29)
Turkey	0.71	0.16 (0.36)	0.66 (0.21)***	0.47 (0.05)**
United Kingdom	0.49	0.16 (0.34)	0.85 (0.10)***	0.40 (0.07)***
United States	0.38	0.33 (0.11)	1.18 (0.02)**	0.08 (0.66)

Notes: (1) All the models are estimated including the intercept (constant) term. The estimates of the intercept term are not reported to conserve space; (2) *, ** and *** indicate the statistical significance and implied rejection of the null hypothesis at 1%, 5% and 10% levels, respectively. Some of the p-values are on the border line of critical region and, thus, strictly do not reject the null hypothesis at the indicated level of significance.

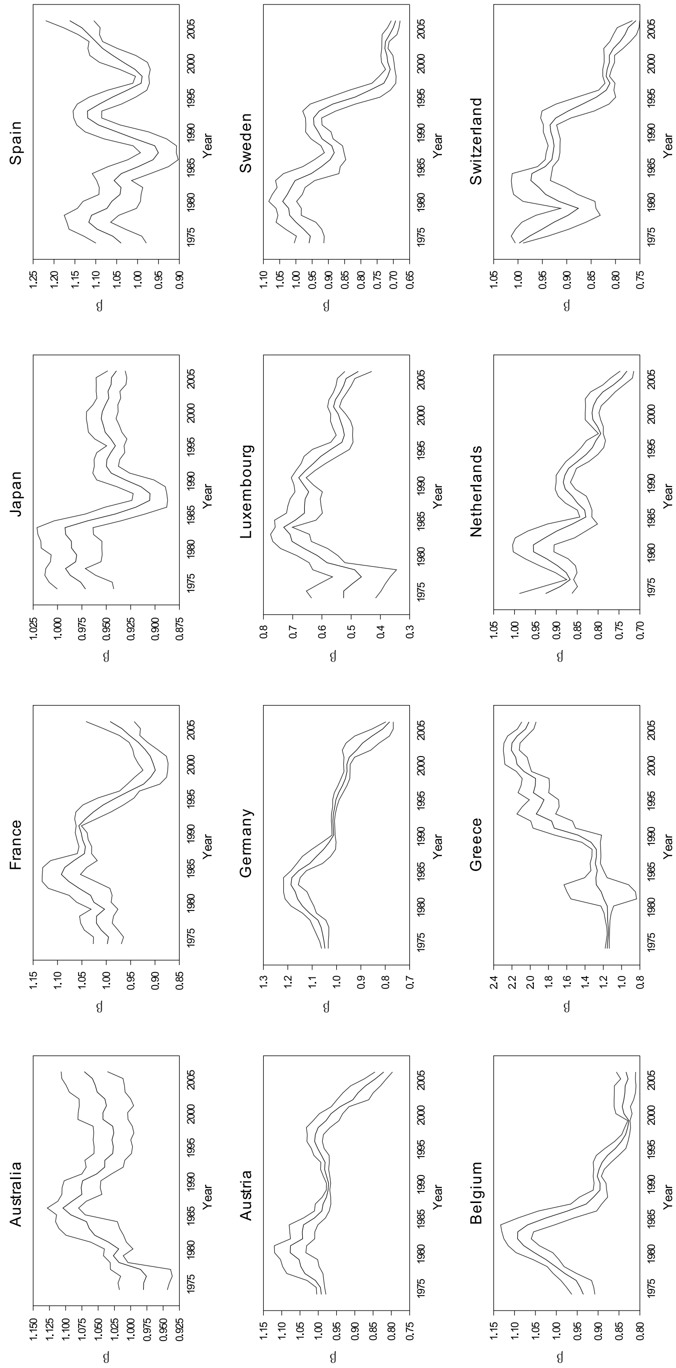
both intercept and slope dummies (entire coefficient vector) to determine the regime shift (Model III).

(8) Model I: Constant; Level Shift: $y_t = \alpha_0 + \alpha_1\delta_t + \beta x_t + \varepsilon_t$

(9) Model II: Constant and Trend; Level Shift with trend: $y_t = \alpha_0 + \alpha_1\delta_t + \phi t + \beta x_t + \varepsilon_t$

(10) Model III: Constant and Slope; Regime Shift: $y_t = \alpha_0 + \alpha_1\delta_t + \beta x_t + \xi\delta_t x_t + \varepsilon_t$

$$\delta_t = \begin{cases} 1; \text{If } t \leq \{\tau T\} \\ 0; \text{If } t > \{\tau T\} \end{cases}; \quad t \in \{1, \ldots, T\}$$


Australia
France
Japan
Spain
Austria
Germany
Luxembourg
Sweden
Belgium
Greece
Netherlands
Switzerland

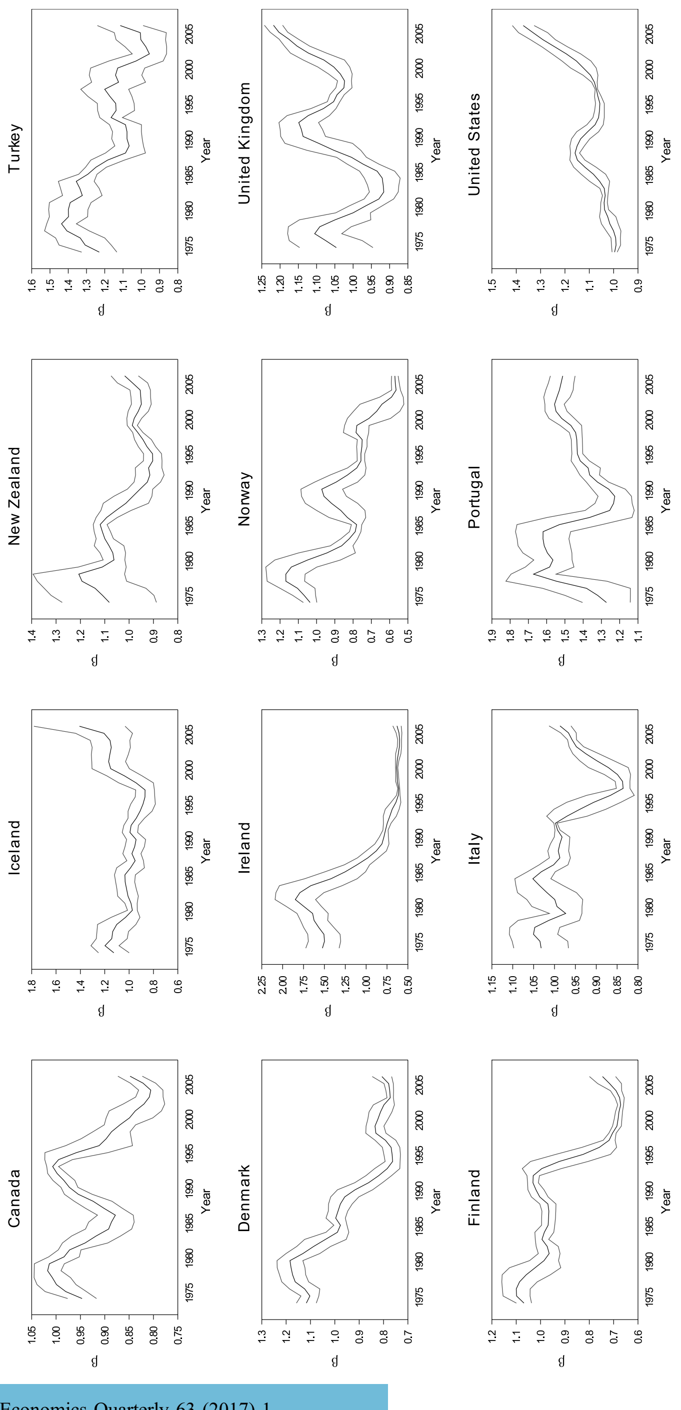


Figure 1: Rolling Regression and the Stability of the Model
[Five Year Moving Window]

The δ_t is a dummy variable that takes value 1 if it is below and value 0 if it is above the unknown break-point, and $\{\cdot\}$ is the integer part. The unknown regime shift parameter, $\tau \in \{0, 1\}$, shows the (relative) timing of the change point in terms of a fraction of the sample space, T. The structural change is reflected in the changes in intercept and/or slope parameters. The second step in OLSGH involves the use of ADF unit root tests on ε_t to test H_0:ε_t~I(1) (no cointegration among I(1) variables) against H_1:ε_t~I(0) (cointegration among I(1) variables) with a single unknown structural break. The observations are trimmed at both beginning and end of the sample space. The ADF(τ) statistics [denoted as GH-ADF] are computed and the grid-search is performed over the trimmed interval to endogenously determine the break-point $\tau \in \{0.15 \times T, 0.85 \times T\}$. The results suggest that the minimized GH-ADF statistic rejects the null hypothesis of no cointegration for (i) Australia (at 5% level), Belgium (at 5% level), Japan (at 10% level), Portugal (at 10% level), Sweden (at 10% level) and Turkey (at 5% level) in the model with intercept dummy (Model I), (ii) Australia (at 10% level), Belgium (at 10% level), Italy (at 10% level), Japan (at 5% level), Switzerland (at 5% level) and Turkey (at 5% level) in the model with intercept and trend dummies (Model II), and (iii) Australia (at 5% level) and Turkey (at 1% level) in the model with intercept and slope dummies (Model III) (Table 9). The OLSGH estimates do not reject the null hypothesis of no cointegration for the remaining countries in the sample.

3.3.3 Tests for Multiple Structural Breaks

The OLSGH assumes a single structural break and precludes the possibility of multiple breaks in the model parameters. Bai and Perron (BP) (1998, 2003) consider linear model and use the dynamic programming algorithm to determine m number of unknown structural breaks and implied m+1 number of regimes. The BP statistics are the generalization of the single-break test statistics of Andrews (1993, 2003) and are robust to the serial-correlation and heterogeneity of residuals under the null hypothesis. Kejriwal and Perron (KP) (2008, 2010) allow I(1) as well as I(0) regressors in the cointegrating model. They derive the limiting distribution of the Sup-Wald test under mild conditions on the errors and regressors for a variety of testing problems. Kejriwal and Perron (2008) show that if the coefficients of the integrated regressors are allowed to change, then the estimated break fractions are asymptotically dependent so that the confidence intervals need to be constructed jointly. If, however, only the intercept and/or the coefficients of the stationary regressors are allowed to change, then the estimates of the break dates are asymptotically independent as in the stationary case analysed by Bai and Perron (1998, 2003). The structural breaks can take place in the form of changes in either intercept or slope parameter of the cointegrating vector. Kejriwal and Perron (2008, 2010) suggest the use of linear DOLS estimator to resolve the endogeneity of regressors and serial-correlation of residuals. The KP results are valid, under very weak conditions, when the potential endogeneity of non-stationary regressors is accounted for via an increasing sequence of

Table 9

OLSGH Test for Cointegration with a Single Structural Break

Country	Minimized GH–ADF Statistics			Break Years		
	Model I	Model II	Model III	Model I	Model II	Model III
Australia	–4.63**	–4.94***	–5.18**	1976	1991	1991
Austria	–1.94	–2.59	–2.30	2001	1991	1989
Belgium	–4.66**	–4.91***	–4.58	1983	1982	1983
Canada	–3.69	–3.61	–3.70	1981	1981	1981
Denmark	–3.91	–3.44	–3.77	1979	1979	1979
Finland	–3.10	–3.86	–3.56	1996	1991	1996
France	–3.43	–3.69	–3.64	1992	1992	1992
Germany	–2.61	–3.43	–2.55	1998	2001	1998
Greece	–3.66	–4.56	–3.92	1988	1988	1989
Iceland	–3.09	–2.74	–4.42	1999	1999	1994
Ireland	–3.78	–4.24	–4.00	1984	1984	1986
Italy	–3.96	–4.80***	–4.56	1984	2001	1991
Japan	–4.53***	–5.29**	–4.42	1981	1984	1981
Luxembourg	–3.38	–4.41	–3.77	1979	1979	1982
Netherlands	–4.02	–3.98	–4.01	2001	1987	2001
New Zealand	–3.84	–4.11	–3.82	1987	1987	1976
Norway	–2.62	–3.09	–3.18	1980	1979	1988
Portugal	–4.53***	–4.61	–4.53	1986	1986	1986
Spain	–4.19	–4.39	–4.43	2001	1981	1995
Sweden	–4.39***	–4.53	–4.39	1994	1994	1992
Switzerland	–3.29	–5.33**	–3.41	1995	1978	1995
Turkey	–4.89**	–5.27**	–5.74*	1999	1999	1999
United Kingdom	–4.08	–3.99	–3.71	1977	1978	1978
United States	–3.03	–3.46	–2.85	2001	1989	2001

Note: *, ** and *** indicate the statistical significance at 1%, 5% and 10% levels, respectively.

lags and leads of the first-differenced dynamic regressors in DOLS. They show that the limiting distributions of the tests based on DOLS are the same as those obtained with static regression under the strict exogeneity.

Both BP and KP suggest three tests for testing multiple structural breaks. The first test is the Sup-Wald test for the null hypothesis of no structural break $(m = 0)$ against the alternative hypothesis of $m = L$ number of arbitrarily fixed structural breaks. The second test is the double maximum (UDmax) test for the null hypothesis of no structural break $(m = 0)$ against the alternative hypothesis of an unknown number of structural breaks between 1 and some upper bound M, such that $1 \leq m \leq M$. The UDmax statistic weighted with marginal p-values across structural breaks becomes the WDmax statistic. The third test is the sequential, $SEQ_T(L+1)/L$, test and it sequentially tests the null hypothesis of L against the alternative hypothesis of L+1 number of structural breaks. The Sup-F(1|0) test is first used to test the null hypothesis of zero versus one break; if the rejection occurs, then the Sup-F(2|1) is used to test the null hypothesis of one versus two breaks; if the rejection again occurs, then the Sup-F(3|2) is used to test the null hypothesis of two versus three breaks, and so on until the non-rejection occurs. The number of breaks are estimated as the number of rejections of the null hypothesis. The model with L number of structural breaks is obtained by global minimization of the residual sum of squares. An alternative to using sequential, $SEQ_T(L+1)/L$, test is to use the Bayesian information criterion (BIC) of Yao (1988) or the modified Schwarz information criterion (LWZ) of Liu et al. (1997) to determine the optimal number of structural breaks. The sequential procedure, however, performs better as compared to BIC and LWZ in that the sequential procedure can easily allow and take into account the effects of possible serial-correlation in errors (Bai and Perron 1998, 2003).

The study estimates the DOLS model (5) using the lags-leads structure of k = {–2, 0, +2} for the first-differenced I(0) regressors for all the countries. Both intercept and slope parameters are allowed to change across regimes. The coefficients of the lagged, contemporaneous and lead I(0) regressors are not allowed to break and, thus, are considered fixed and invariant over time. The inclusion of I(0) regressors whose coefficients are not allowed to change does not alter the limit distribution. The DOLS estimation is also carried out using one lower, k = {–1, 0, +1}, and one higher, k = {–3, 0, +3}, lags-leads structures of the I(0) regressors to assess the robustness of results across model structures. The results obtained from the alternative model structures were generally consistent in terms of the number and locations of the break-points and are, therefore, reported only for the model with k = {–2, 0, +2}[5]. These results suggest that the F test rejects the null hypothesis of no structural break (m = 0) against the alternative hypothesis of m = L number of breaks for all the dates and across all the sample countries (Table 10). The Sup-F(m) statistics are significant for all the countries, with L running between 1 and 5. This suggests that at least one structural break could be present in the relationship between the model series. The UDmax statistics are highly significant at 1% level for most countries, reinforcing the presence of at least one structural break in the model. The sequential F test is finally used to test the null hypothesis of L against the alternative hypothesis of L+1 number of structural breaks, and determine the optimal number of break-points. The $SEQ_T(L+1)/L$ statistics are significant at 1% level for most countries (Table 10). The test rejects the null hypothesis of L=4 breaks and suggests the presence of five breaks and implied six number of regimes for all the countries. The estimates for the number of structural breaks obtained from the sequential test are at variance with the estimates suggested by the BIC and LWZ statistics. The break years show only minor variations across the sample countries.

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[5] The results from the DOLS estimations, carried out using the alternative model structures of k = {–1, 0, +1} and k = {–3, 0, +3} for all the countries, are available from the author on request.

Table 10

Sup-F Test for Zero versus an Unknown Number of Structural Breaks

Country	Sup-F(m) Statistics [H_0: m = 0; H_1: m = L]					UDmax(L)
	Sup-F(1|0)	Sup-F(2|0)	Sup-F(3|0)	Sup-F(4|0)	Sup-F(5|0)	
Australia	6.87	41.74*	8.15	95.45*	31.43*	95.45* (4)
Austria	13.51***	20.73*	21.86*	26.30*	20.87*	26.30* (4)
Belgium	57.18*	34.09*	32.72*	13.92*	39.37*	57.18* (5)
Canada	8.54	48.23*	39.44*	68.28*	53.99*	68.28* (4)
Denmark	23.96*	39.80*	35.17*	83.05*	16.74*	83.05* (4)
Finland	7.46	101.15*	103.61*	242.48*	234.79*	242.48* (4)
France	42.66*	201.05*	28.28*	138.40*	175.22*	201.05* (2)
Germany	24.06*	28.72*	9.05	13.13*	10.61*	28.72* (2)
Greece	6.13	19.35*	134.67*	154.75*	18.29*	154.75* (4)
Iceland	15.89**	13.93**	9.77	7.64	3.98	15.89** (1)
Ireland	89.98*	50.33*	111.00*	53.05*	196.34*	196.34* (5)
Italy	0.38	40.64*	28.18*	101.59*	29.65*	101.59* (4)
Japan	7.89	17.30*	11.37	46.02*	76.57*	76.57* (5)
Luxembourg	1.13	10.88	27.26*	14.51*	22.55*	27.26* (3)
Netherlands	4.22	9.66	48.34*	16.93*	25.83*	48.34* (3)
New Zealand	75.11*	39.68*	78.66*	63.77*	61.25*	78.66* (3)
Norway	30.25*	18.05*	38.06*	56.40*	33.04*	56.40* (4)
Portugal	38.50*	56.66*	52.42*	42.31*	43.61*	56.66* (2)
Spain	9.52	13.24**	10.09	18.43*	9.54*	18.43* (4)
Sweden	77.27*	63.99*	92.23*	146.54*	307.83*	307.83* (5)
Switzerland	32.34*	11.06***	104.53*	206.97*	114.01*	206.97* (4)
Turkey	6.96	5.28	13.92*	14.81*	5.31	14.81** (4)
United Kingdom	6.29	14.79*	33.46*	20.01*	13.31*	33.46* (3)
United States	15.64*	7.30	7.62	14.00*	9.50*	15.64** (1)
Significance Level	Critical values					
1%	17.67	14.73	12.21	10.77	8.82	17.67
5%	14.30	12.11	10.41	9.19	7.64	14.47
10%	12.36	11.01	9.60	8.45	6.96	12.64

Notes: (1) *, ** and *** indicate the statistical significance at 1%, 5% and 10% levels, respectively; (2) The critical values are taken from Kejriwal and Perron (2010), Table 1a.

Table 11

Sequential, $SEQ_T(L+1)/L$, Test for L versus (L + 1) Number of Breaks and the Minimized BIC and LWZ Statistics

	$SEQ_T(L+1)/L$					Number of Breaks		
	[H_0: L breaks; H_1: L + 1 breaks]					Sequential F Test	Minimized BIC and LWZ	
Country	Sup–F(1|0)	Sup–F(2|1)	Sup–F(3|2)	Sup–F(4|3)	Sup–F(5|4)	Breaks	BIC(L)	LWZ(L)
Australia	6.87	83.48*	24.46*	381.79*	157.15*	5	–12.66 (4)	–11.04 (4)
Austria	13.51	41.46*	65.58*	105.21*	104.34*	5	–11.10 (5)	–9.41 (4)
Belgium	57.18*	68.17*	98.15*	55.66*	196.87*	5	–11.49 (5)	–9.51 (5)
Canada	8.54	96.45*	118.32*	273.11*	269.96*	5	–11.14 (5)	–9.44 (4)
Denmark	23.96*	79.61*	105.50*	332.20*	83.69*	5	–11.11 (4)	–9.49 (4)
Finland	7.46	202.30*	310.84*	969.92*	1173.93*	5	–10.80 (5)	–8.90 (4)
France	42.66*	402.10*	84.85*	553.61*	876.10*	5	–12.38 (5)	–10.44 (2)
Germany	24.06*	57.45*	27.15*	52.54*	53.07*	5	–11.03 (2)	–9.91 (2)
Greece	6.13	38.71*	404.00*	619.01*	91.45*	5	–11.54 (4)	–9.93 (4)
Iceland	15.89**	27.85*	29.32*	30.57*	19.88**	5	–8.69 (2)	–7.58 (2)
Ireland	89.98*	100.66*	332.99*	212.20*	981.70*	5	–10.44 (5)	–8.45 (5)
Italy	0.38	81.27*	84.55*	406.35*	148.26*	5	–11.67 (4)	–10.05 (4)
Japan	7.89	34.61*	34.11*	184.06*	382.84*	5	–13.01 (5)	–11.03 (5)
Luxembourg	1.13	21.76*	81.78*	58.05*	112.73*	5	–10.70 (5)	–8.97 (3)
Netherlands	4.22	19.32**	145.03*	67.73*	129.17*	5	–11.86 (3)	–10.51 (3)
New Zealand	75.11*	79.35*	235.97*	255.09*	306.23*	5	–10.33 (5)	–8.64 (3)
Norway	30.25*	36.10*	114.18*	225.62*	165.21*	5	–8.53 (4)	–6.91 (4)
Portugal	38.5*	113.32*	157.27*	169.23*	218.06*	5	–10.14 (5)	–8.39 (3)

Continued next page

Table 11 (continued)

	$SEQ_T(L+1)/L$					Number of Breaks		
	[H_0: L breaks; H_1: L + 1 breaks]					Sequential F Test	Minimized BIC and LWZ	
Country	Sup–F(1|0)	Sup–F(2|1)	Sup–F(3|2)	Sup–F(4|3)	Sup–F(5|4)	Breaks	BIC(L)	LWZ(L)
Spain	9.52	26.47*	30.28*	73.71*	47.69*	5	–11.42 (4)	–9.80 (4)
Sweden	77.27*	127.98*	276.69*	586.17*	1539.17*	5	–12.55 (5)	–10.57 (5)
Switzerland	32.34*	22.12*	313.59*	827.89*	570.07*	5	–11.63 (4)	–10.01 (4)
Turkey	6.96	10.57	41.75*	59.25*	26.54*	5	–10.02 (4)	–8.41 (4)
United Kingdom	6.29	29.59*	100.38*	80.03*	66.57*	5	–10.56 (3)	–9.22 (3)
United States	15.64***	14.59	22.86*	56.02*	47.50*	5	–11.94 (4)	–10.32 (4)
Significance Level	Critical values							
1%	19.04	19.35	19.90	19.99	20.01			
5%	15.65	16.61	17.12	17.66	17.85			
10%	14.26	15.02	15.64	16.02	16.51			

Notes: (1) *, ** and *** indicate the statistical significance at 1%, 5% and 10% levels, respectively; (2) The critical values are taken from Kejriwal and Perron (2010), Table 3a; (3) The figures in parenthesis associated with the minimized values of the BIC and LWZ statistics are the L number of breaks.

Table 12

Estimated Break-Points and the 95 % Confidence Intervals

Country	Break–Points and the 95 % Lower and Upper Confidence Bands				
	m=1	m=2	m=3	m=4	m=5
Australia	1980 [1979 – 1981]	1984 [1983 – 1985]	1988 [1987 – 1989]	1992 [1991 – 1993]	1997 [1996 – 1998]
Austria	1978 [1977 – 1979]	1983 [1982 – 1984]	1989 [1988 – 1990]	1993 [1992 – 1994]	1997 [1996 – 1998]
Belgium	1978 [1977 – 1979]	1983 [1982 – 1984]	1988 [1987 – 1989]	1992 [1991 – 1993]	1997 [1996 – 1998]
Canada	1978 [1977 – 1978]	1982 [1981 – 1983]	1986 [1985 – 1987]	1990 [1989 – 1991]	1997 [1996 – 1998]
Denmark	1979 [1977 – 1980]	1983 [1982 – 1984]	1989 [1988 – 1990]	1993 [1992 – 1994]	1997 [1996 – 1998]
Finland	1979 [1978 – 1980]	1983 [1982 – 1984]	1987 [1986 – 1988]	1991 [1990 – 1992]	1995 [1994 – 1996]
France	1978 [1977 – 1979]	1983 [1982 – 1984]	1988 [1987 – 1989]	1992 [1991 – 1993]	1997 [1996 – 1998]
Germany	1979 [1978 – 1980]	1983 [1982 – 1984]	1987 [1986 – 1988]	1991 [1989 – 1992]	1996 [1995 – 1997]
Greece	1981 [1980 – 1982]	1985 [1984 – 1985]	1989 [1988 – 1990]	1993 [1992 – 1994]	1997 [1996 – 1998]
Iceland	1978 [1977 – 1979]	1983 [1982 – 1984]	1987 [1985 – 1988]	1991 [1990 – 1992]	1996 [1995 – 1997]
Ireland	1978 [1977 – 1979]	1982 [1981 – 1983]	1989 [1988 – 1990]	1993 [1992 – 1994]	1997 [1996 – 1998]
Italy	1978 [1977 – 1979]	1983 [1982 – 1983]	1987 [1986 – 1989]	1991 [1990 – 1992]	1995 [1994 – 1996]
Japan	1978 [1977 – 1979]	1982 [1981 – 1983]	1987 [1986 – 1988]	1992 [1991 – 1993]	1997 [1996 – 1998]
Luxembourg	1978 [1977 – 1979]	1983 [1982 – 1984]	1988 [1987 – 1989]	1992 [1991 – 1993]	1997 [1996 – 1998]
Netherlands	1979 [1978 – 1980]	1983 [1982 – 1984]	1988 [1987 – 1989]	1993 [1992 – 1994]	1997 [1996 – 1998]
New Zealand	1979 [1978 – 1980]	1983 [1982 – 1984]	1988 [1987 – 1989]	1993 [1992 – 1994]	1997 [1996 – 1998]
Norway	1978 [1977 – 1979]	1982 [1981 – 1983]	1987 [1986 – 1988]	1992 [1992 – 1993]	1996 [1995 – 1997]
Portugal	1978 [1977 – 1979]	1984 [1983 – 1985]	1988 [1987 – 1988]	1992 [1991 – 1993]	1996 [1995 – 1997]
Spain	1978 [1977 – 1979]	1985 [1980 – 1989]	1989 [1988 – 1990]	1993 [1992 – 1994]	1997 [1996 – 1998]
Sweden	1978 [1977 – 1979]	1982 [1981 – 1983]	1988 [1987 – 1989]	1993 [1992 – 1994]	1997 [1996 – 1997]
Switzerland	1981 [1980 – 1982]	1985 [1984 – 1986]	1989 [1988 – 1990]	1993 [1992 – 1994]	1997 [1996 – 1998]
Turkey	1979 [1978 – 1980]	1983 [1982 – 1984]	1987 [1986 – 1988]	1992 [1991 – 1993]	1996 [1995 – 1997]
United Kingdom	1978 [1977 – 1979]	1982 [1981 – 1983]	1986 [1985 – 1987]	1991 [1990 – 1992]	1997 [1996 – 1998]
United States	1978 [1977 – 1979]	1982 [1981 – 1983]	1988 [1987 – 1989]	1993 [1991 – 1994]	1997 [1996 – 1998]

3.3.4 Test for Cointegration with Multiple Structural Breaks

The single-equation OLSGH tests the null hypothesis of no cointegration in the presence of a one-time break in either intercept or both intercept and slope parameters. Kejriwal (2008) extends the one-time structural break test of Arai and Kurozumi (2005) and develops the test for the null hypothesis of cointegration in the presence of multiple breaks in a single-equation setting. While the single-equation methodology is useful, it imposes, at the most, a single cointegrating vector on the model series and, thus, precludes the possibility of more than one equilibrium relationship among the model variables. Johansen, Mosconi and Nielsen (JMN) (2000) extend the ML cointegration estimator of Johansen (1988, 1995) and test the null hypothesis of no cointegration in the presence of exogenously determined multiple structural breaks in the equilibrium relationship in a system setting. The asymptotic distribution of the rank test and the critical values for the model parameters are simulated using the response surface approximation. The JMN estimator is useful to determine the number of cointegrating vectors in a multivariate system in the presence of structural breaks at the known points in time.

The study uses the JMN test and tests the null hypothesis of no cointegration against the alternative hypothesis of cointegration in the presence of multiple (five) structural breaks. The break years suggested by the sequential, $SEQ_T(L+1)/L$, test of Kejriwal and Perron (2008, 2010) are used to set the exogenous break dummies and estimate the VAR model. The results suggest that the long-run relationship between the model series holds even after accounting for structural breaks in level (Table 13). The coefficients of the break dummies are not reported for brevity[6]. The long-run slope parameter of saving, β, is counter-intuitively either unduly large (more than unity) or unduly small (negative) with little economic interpretation for a number of countries. The slope parameter, however, carries the expected signs and meaningful magnitudes for some of the countries in the sample.

4. Financial Autarky Versus Financial Globalization: Some Analytics

The above stylized evidence suggests that the degree of IMC and the integration of financial markets vary across countries. The support for heterogeneous degrees of integration of financial markets across countries has important implications for the financial architecture of the global economy in general and for the macroeconomic stabilization policies of the sample countries in particular. The investment in the countries with moderate mobility of capital is financed by both domestic and foreign saving. The reliance on domestic saving in the countries with low to moder-

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6 The complete set of results for the cointegration test of Johansen, Mosconi and Nielsen (2000), including the coefficients of the break dummies, is available from the author on request.

Table 13

ML System Estimates of the Long-Run Model with Structural Breaks [VAR lag k =2]

Country	β	Structural Break Years	Eigenvalues		λ-trace Test		λ-trace Test@	
			H_0: r = 0	H_0: r ≤ 1	H_0: r = 0	H_0: r ≤ 1	H_0: r = 0	H_0: r ≤ 1
Australia	0.78	1980 1984 1988 1992 1997	0.72	0.31	58.15**	13.17**	54.46**	12.19**
Austria	0.06	1978 1983 1989 1993 1997	0.47	0.25	32.33**	10.06**	29.72**	8.12**
Belgium	0.53	1978 1983 1988 1992 1997	0.56	0.32	42.20**	13.23**	39.24**	12.57**
Canada	–2.98	1978 1982 1986 1990 1997	0.59	0.41	49.44**	18.39**	44.49**	17.95**
Denmark	0.38	1979 1983 1989 1993 1997	0.60	0.39	49.26**	17.13**	44.20**	16.10**
Finland	–0.24	1979 1983 1987 1991 1995	0.49	0.19	30.66**	7.17**	28.05**	7.05**
France	–0.29	1978 1983 1988 1992 1997	0.55	0.40	45.91**	17.82**	41.15**	16.28**
Germany	–1.04	1979 1983 1987 1991 1996	0.50	0.19	31.20**	7.30**	27.27**	6.50**
Greece	1.03	1981 1985 1989 1993 1997	0.72	0.27	55.58**	10.90**	51.77**	10.63**
Iceland	–4.69	1978 1983 1987 1991 1996	0.64	0.16	42.03**	5.98**	38.61**	5.73**
Ireland	0.45	1978 1982 1989 1993 1997	0.57	0.50	53.91**	24.11**	49.34**	22.78**
Italy	–0.72	1978 1983 1987 1991 1995	0.52	0.42	44.62**	18.80**	40.66**	17.76**
Japan	0.87	1978 1982 1987 1992 1997	0.74	0.53	73.14**	26.38**	64.90**	24.98**
Luxembourg	1.45	1978 1983 1988 1992 1997	0.62	0.39	51.26**	17.54**	48.22**	13.86**
Netherlands	–6.83	1979 1983 1988 1993 1997	0.39	0.37	33.86**	16.32**	30.04**	14.21**
New Zealand	3.73	1979 1983 1988 1993 1997	0.57	0.25	39.69**	10.17**	37.18**	NA
Norway	–1.87	1978 1982 1987 1992 1996	0.45	0.25	30.74**	9.86**	27.46**	9.35**
Portugal	–0.13	1978 1984 1988 1992 1996	0.47	0.34	36.16**	14.29**	33.16**	13.93**
Spain	2.73	1978 1985 1989 1993 1997	0.54	0.40	44.94**	17.71**	39.92**	16.52**
Sweden	–0.11	1978 1982 1988 1993 1997	0.60	0.45	52.99**	21.18**	47.52**	20.98**
Switzerland	0.51	1981 1985 1989 1993 1997	0.42	0.31	32.10**	12.76**	28.18**	12.31**
Turkey	6.73	1979 1983 1987 1992 1996	0.57	0.38	46.04**	16.77**	43.07**	16.50**
United Kingdom	0.06	1978 1982 1986 1991 1997	0.55	0.37	44.20**	16.20**	40.41**	15.63**
United States	0.29	1978 1982 1988 1993 1997	0.64	0.44	56.14**	20.48**	49.66**	18.80**

Notes: (1) The r denotes the number of cointegrating vectors; (2) @ indicates the λ-trace corrected for small sample; (3) The 95% critical values are 15.408 for H_0: r = 0 and 3.841 for H_0: r ≤ 1; (4) ** denotes the statistical significance at 5% level; (5) $ Long-run coefficients of the first cointegrating vector are normalised on I/Y; (6) NA implies that the value could not be computed.

ate mobility of capital underlines the need for the acceleration of domestic saving to finance the accumulation of capital and keep the current account imbalances in sustainable bounds. The investment in the countries with high IMC is financed significantly by a world pool of capital, and the domestic saving flows in response to higher returns (relative to risks) available in the international financial markets. The access to international financial markets (for lending) tends to impinge upon the efficacy of domestic economic policies in accelerating domestic investment through the increase in domestic saving. This section undertakes a systematic account of the policy implications of SI correlations for the mobility of capital and integration of financial markets across countries.

4.1 Speculative Capital Flows and the Financial Globalization

The financial globalization in a perfectly competitive and symmetric information environment facilitates the diversification of investment portfolios across international financial assets, global sharing of financial risks, maximization of risk-adjusted rates of returns, and the efficient allocation of world capital resources. The financial integration imposes greater discipline on the fiscal and monetary policies. The Mundell-Fleming model postulates the perfect mobility of capital and frictionless integration of international financial markets. It views capital flows as sensitive to the interest-rate differentials across countries. Any disparity between domestic and foreign interest rates is offset by the inflow (outflow) of capital and is reflected in (i) the accumulation (depletion) of foreign exchange reserves and implied balance of payments surpluses (deficits) under the fixed exchange rate regime and (ii) an appreciation (depreciation) in the external value of domestic currency under the floating exchange rate regime. The well-documented gains of financial openness, however, remain surrounded by the concerns for the destabilizing effects of the high-resolution and speculative capital flows. The major concern in the countries with high IMC is the vulnerability to the shocks (systematic or stochastic) of speculations by the international investors. The episodes of financial crises witness that the short-term and excessively volatile capital flows could destabilize even the informationally efficient and financially solvent systems (with assets more than liabilities), and make the economies prone to financial calamities with self-fulfilling runs on currencies and panic reversals of capital inflows. The theoretical models commonly postulate that the investors are rational, well-informed, farsighted, and forward-looking. The financial crises could arise from the irrational speculations and over-reactions of investors compared to the theoretically postulated benchmark norms. While the financial markets could simply be the belated messengers, rather than causes of financial crises, the observed episodes of financial crises accentuate the concerns for the risks involved in investing in foreign financial markets. This seems to be one of the factors catalytic to the moderate, rather than high, mobility of capital observed across most countries.

The foreign borrowings and excessive accumulations of external debts increase the risk-premium and potentially lead to the generation of the 'vicious circle' of 'high borrowings and debt defaults'. The high risk premium and the vulnerability to the 'speculative attacks' on the external value of domestic currency induce the market-driven need to reduce external borrowing and, thus, make the foreign debt a self-limiting and finite phenomenon. The market-driven limits on external borrowing and the accumulations of foreign debt are commonly complemented by the policy-determined impositions of capital controls so as to reduce the high-resolution and short-term capital flows and minimize the amplitudes of volatility in the financial markets. The foreign exchange reserves provide a buffer against capital outflows apart from financing imports. This is plausibly one of the reasons that the central banks world over have not relinquished their intervention in the foreign exchange markets and the accumulation of foreign exchange reserves even in the regime of floating exchange rate with increased mobility of capital and integration of financial markets.

The FH sample (1960–74) pertains to the Bretton Woods regime that prevailed until the early-1970s, and the then high SI correlations could be a corollary of the capital controls imposed to support the fixed exchange rates and independent monetary policies. The policy trilemma of maintaining (i) fixed exchange rate system, (ii) autonomous monetary policy and (iii) open capital account narrow down the policy choices to, at the most, two (any two) of the three policy objectives. If a country chooses fixed exchange rate regime and free IMC, then it has to abandon its monetary autonomy; if a country chooses monetary autonomy and IMC, then it has to adopt flexible exchange rate system; and if a country combines fixed exchange rate system with monetary autonomy (at least in the short run), then it would need to restrict IMC. The Keynesian perception that the economy is inherently unstable necessitates the need for the intervention of short-term management (monetary and fiscal) policies to stabilize the economy around the long-run natural rate level of output (employment). The need for having autonomous monetary policy reduces the option to choosing between fixed exchange rate system and open capital account. Besides, the countries commonly prefer to have a mid-way path and choose neither completely 'fixed' nor 'flexible', but a moderately flexible exchange rate system. The exceptions are the countries with currency boards or some such fixed exchange rate arrangements. Thus, the preponderant preferences for independent monetary policy and moderately (rather than perfectly) flexible exchange rate system effectively reduce the policy choice to having a modestly open capital account with moderate mobility of capital, rather than completely open capital account with perfectly free mobility of capital. The study finds that the mobility of capital varies across countries. The imperfect integration of international goods and services markets tends to be accompanied by the imperfect integration of international financial markets. The countries forming the trading blocs and entering into the free (preferential or regional) trade agreements commonly trade more with their 'natural' trading partners and, thus, have bloc-local (within the trading bloc) high mobility of capital as compared to the trade (and mobility of capital) outside the trading bloc and across the globe (see Singh 2010). The Maastricht Treaty (February 1992) and the formation of the European Union plausibly contributed to the removal of policy restrictions on capital accounts and, thus, to the integration of financial markets among the member countries.

4.2 Informational Inefficiency in International Financial Markets

The evidence on the informational efficiency in the financial markets remains mixed and inconclusive. The financial markets, and more particularly the international financial markets, are characterized by asymmetric information problem. The distributions of the asset portfolios of investors remain skewed towards the holdings of domestic, rather than foreign, financial assets, despite the well-documented gains of international diversification and financial openness of the economy. The majority of investors remain parochial in outlook and choose to retain most of their wealth at home, mainly due to the information asymmetries in the international financial markets. The asymmetries arising from the differences in the quantity and quality of information between domestic and foreign investors, uncertainties involved in investing in unknown international financial markets, and several unfamiliarities with foreign products, business practices, firms, accounting standards, and regulatory environments accentuate home-bias in the asset portfolios of investors and impede the perfect integration of international financial markets. Even within the domestic economy, the saving portfolios of households remain dominated by the holdings of the safe and risk-free vector of financial assets. The investment in equities tends to remain relatively small, reflecting the 'equity premium puzzle' in the capital markets. The asymmetries in information and the risks arising from moral hazards are the major impediments to the household investment in equities. The 'lemon' firms commonly offer high-yield (but high-risk) securities compared to the high-grade 'peach' firms that commonly offer low-yield (but low-risk and investment-grade) securities. It is difficult for a median investor to afford costly information and discern between 'lemons' and 'peaches' in the capital market. If a risk-averse household is hesitant to investing in domestic debt and equity securities markets, it is consistently likely to be at least equally hesitant to investing in still more unfamiliar foreign securities (debt and equity) markets.

It could reasonably be argued that the foreign investors may overcome the 'lemons problem' by investing in a diversified portfolio of publicly traded stocks of domestic firms. But the set of firms that list their shares on public stock exchanges may not be representative of all the domestic firms. The lemons may endeavour to list their shares in the hope that the market will overvalue them (Lewis 1999). Besides, even when buying diversified portfolios, foreign investors may be still at informational disadvantage given that the information available to domestic investors is not fully conveyed through the market prices, due to the noise in these prices (Gordon and Bovenberg 1996). The price system, in general, does not reveal all the information about 'the true value' of the risky asset (Grossman and Stiglitz 1980). The imperfect pass-through of information to asset prices and several informational inefficiencies in the financial markets impede the mobility of capital across countries. The development of information technology has indeed helped reduce (though not completely remove) 'lemons problem' and that financial innovations and the resultant development of risk-management financial derivative products (futures, forward, options and swaps) have helped hedge (though not perfectly hedge) the exposures to exchange rate risks. The fixed exchange rate system minimizes the risks arising from the frequent depreciations (appreciations) of the external value of domestic currency. This advantage lost with the adoption of floating exchange rate system since the early-1970s has partially been compensated by the advent and development of information technology and the innovations of financial derivative products. Obstfeld (1998) observes that the worldwide trends in financial opening in the 1990s have restored a degree of IMC not seen since the beginning of this century. The integration of international financial markets has increased in both developed and developing countries during the last three decades due to financial innovations and liberalization, technological breakthrough and the growth of world trade. Nevertheless, the investment preferences still seem to lean towards the holdings of the domestic financial assets. The home-bias tends to dominate the foreign bias in the asset portfolio of investors, notwithstanding the increase in the mobility of capital across countries.

4.3 Bounded Rationality

The informational inefficiency in the international financial markets combined with the 'bounded rationality' of investors tends to impinge upon the optimal decision-making. The developments in behavioural economics (finance) and the evolution of neuroeconomics (neurofinance) have led to the reevaluation of the conventional concept of rationality used in economics. Conlisk (1996) discusses four reasons for incorporating bounded rationality in economic models. First, there is abundant empirical evidence that it is important. Second, the models of bounded rationality have proved themselves in a wide range of impressive work. Third, the standard justifications for assuming unbounded rationality are unconvincing. Fourth, the deliberation about an economic decision is a costly activity, and good economics requires entertaining all costs (Conlisk 1996). Rieskamp et al. (2006) argue that the bounded rationality is a more accurate description of human behavior taking into account that the people make decisions with limited time, knowledge, and computational power. The financial decision-making depends on a number of factors including the aversion to risk, asymmetry in the profit-and-loss response, asymmetry in the measurement of risk on home versus foreign financial assets, cognitive limitations, noise-trading, emotional reaction and defensive behaviour in response to financial fears, over-assessment of financial risks, over-reaction to new information and economic (policy) changes, limitations of knowledge and computational capacity, and the myopic evaluations of risks / returns.

The rational ignorance or the bounded rationality of investors in conjunction with the imperfect information leads to the sub-optimal decisions with regard to investment in financial assets. The future, which is unknown and uncertain, is central to taking investment decisions. The rational expectation and market efficiency hypotheses suggest that the agents use all the available information to form expectations and make well-informed and rational economic (financial) decisions. The information available to the agents, however, may itself be imperfect in terms of the potential risks and returns on assets. The agents commit systematic and sequential errors in forming expectations and making optimal decisions, due to the lack of perfect knowledge and foresight, non-rational expectations, and 'bounded rationality'. These factors contribute to home-bias in the asset portfolios of investors and impede the mobility of capital across unfamiliar international financial markets. Even the rational agents with perfect information may be unable to predict the outcomes of their investment in uncertain future.

4.4 Intertemporal Budget Constraint and SI Correlations

The intertemporal optimization approach to current account interprets high SI correlation in terms of the validity of intertemporal budget constraint and the sustainability of current account deficits, rather than as an index of capital mobility. It postulates high international mobility of capital analogous to the conventional forerunner, the Mundell-Fleming model. The models developed in intertemporal setting derive the consumption and saving decisions of the forward-looking optimizing agents from the maximization of the utility function subject to the intertemporal budget constraint. Any intertemporal disparity between the paths of consumption and output (saving and investment) is reflected in the accumulation (depletion) of net foreign assets. A country can lend (borrow) resources from the rest of the world and avoid a sharp expansion (contraction) in consumption and investment in case of a temporary excess (short-fall) in income. The trade (current account) deficit (surplus) at any given time is viewed as reflecting the transfer of consumption opportunities across time in the intertemporal optimizing models, rather than reflecting any economic disequilibrium as viewed in the conventional non-optimizing models of trade and current account balance.

The strong diminishing returns at home, weak investment risks abroad, and the diversification benefits induce foreign-bias in the asset portfolios of domestic investors. The marginal unit of wealth arising from a positive income shock is invested in foreign assets and the response of current account is equal to the saving generated by a shock. The saving depends on the intertemporal consumption decisions and that investment flows to where it finds higher real returns. The intertemporal budget constraint may not allow the countries to run high and perpetual current account deficits and the solvency constraint requires the long-run relationship between domestic saving and investment. The SI correlations, as such, tend to be high regardless of the degree of capital mobility and the integration of financial markets across countries. The study finds support for the cointegrating relationship between domestic saving and investment for most countries. The presence of such cointegration points towards the validity of intertemporal budget constraint and the sustainability of current account deficits.

5. Conclusions

This study has examined the long-run relationship between domestic saving and investment and taken a country-by-country account of the mobility of capital and integration of financial markets across countries. The analysis is carried out for a comprehensive set of 24 OECD countries. The OLSEG, OLSGH and ML estimates of the model consistently suggest that the domestic saving and investment are cointegrated for some, but not for all the countries. The degree of capital mobility and integration of financial markets vary across countries. The slope parameter of saving remains well above zero for most countries. The study finds support for structural breaks in the slope parameter of saving for almost all the countries. The JMN test suggests that the cointegrating relationship prevails even after accounting for structural breaks in the model parameters. The support for cointegrating relationship between domestic saving and investment suggests the sustainability of current account deficits and the solvency of intertemporal budget constraint across several countries. An asymmetrically low proportion of investment is financed by domestic saving in the countries with high mobility of capital. The investment in the countries with low to moderate mobility of capital is financed by both domestic and foreign saving. The reliance on domestic saving underlines the need to accelerate saving to finance the accumulation of capital and keep the current account imbalances in sustainable bounds. The investment in the countries with high mobility of capital is financed by a world of pool capital, and the domestic saving flows in response to higher returns available in the international financial markets. The major concerns for the countries with high mobility of capital are the vulnerability to the speculative (systematic or stochastic) expectations (rational or irrational) of international investors, sustainability of current account deficits, adequacy of foreign exchange reserves, and the stability of the financial system.

The switch from fixed to flexible exchange rate regime since the early-1970s provided higher degree of policy freedom, and enabled the dismantling of policy restrictions on capital flows that were placed during the Bretton Woods system. Nevertheless, the approach to the liberalization of capital account has been hesitant and cautious because of the concerns for the destabilizing effects of the high-resolution and speculative capital flows. The preference for ‘moderately’, rather than ‘perfectly’, floating exchange rate regime required a compromise in terms of having ‘moderate’, rather than ‘complete/full’ liberalization of capital account. The mobility of capital has increased overtime, and a part of such increase comes as a corollary of the increased openness to trade in goods and services. Such increased mobility of capital, however, is still far below the benchmark standards postulated in the Mundell-Fleming as well as intertemporal optimizing models. The countries forming the trading blocs and entering into the free (preferential or regional) trade agreements commonly trade more in terms of both 'goods' (trade openness) and 'financial assets' (financial openness) within the trading bloc, The trading outside the bloc, be it in terms of 'goods' or be it in terms of 'financial assets', tends to be relatively less than the trading within the bloc. The countries, thus, commonly have high bloc-local (within the trading bloc) mobility of capital as compared to the trade (and mobility of capital) outside the trading bloc and across the globe. The aversions to the short-term and speculative capital flows and to the excessive accumulations of external debt impinge upon the perfect mobility of capital and frictionless integration of international financial markets.

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