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What do Savings-Investment Correlations tell us about the International Capital Mobility of Less Developed Countries?

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Abstract

This paper investigates the extent of capital mobility with respect to less developed countries over the study period 1979-2001. For this purpose, the Feldstein-Horioka equation linking domestic savings and investment is estimated. However, there is a novel empirical approach in this study based on the employment of panel data methods of cointegration testing and estimation of the long-run saving retention coefficient. There is strong evidence of cointegration between domestic savings and investment. Panel estimation based on fully modified ordinary least squares indicates that capital is imperfectly mobile with a long-run saving retention coefficient of about one-third. This low value suggests that capital mobility, though imperfect, is nonetheless quite high. Further estimation based on sub-groups comprising Asian and Latin American countries points towards a similar degrees of capital mobility.

- **JEL Classifications:** C5, F3, F4, O5
- **Key words:** LDCs, Capital mobility, Savings, Investment, Panel data, Cointegration

I. Introduction

For a variety of reasons, the extent of capital mobility can have profound macro- and microeconomic implications for the well being of both developed and less

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developed countries (LDCs). Measuring capital mobility, however, has proved to be problematic. On the one hand, one might follow Frankel (1992) and use covered interest parity (CIP) as the most appropriate indicator of the degree of financial integration and therefore capital mobility across national boundaries. Essentially, this *direct approach* is based on testing the law of one price in the context of identical financial assets where the price of assets, denominated in different currencies, with similar risk and maturity characteristics tend to equality easily and quickly. For LDCs however, the presence of markets that are illiquid and difficulties in asset comparability contribute towards data limitations that inhibit the formal testing of CIP. An alternative way forward is to consider an *indirect approach* that concentrates on the effects of capital mobility on macroeconomic aggregates such as the relationship between domestic savings and investment [Feldstein and Horioka (1980)]. Here it is argued that there is no need for domestic savings and investment to be correlated with each other if perfect capital mobility enables arbitrage to equalize yields to investors. Domestic savings react to international rates of return and so investment is funded from the world capital market through a current account deficit. If, however, capital were perfectly immobile then one would expect domestic savings and investment to be characterized by a correlation coefficient of unity. Obtaining a correlation of saving and investment close to one in their cross-section analysis for sixteen industrialized OECD countries for the 1960–1974 period, led Feldstein and Horioka (1980) to reject the perfect capital mobility assumption. A number of subsequent analyses using cross-section or times series data have confirmed Feldstein and Horioka's (F-H) results and have attempted to reconcile them with the capital mobility hypothesis (see, *inter alia*, Dooley *et al*. (1987), Vos (1988), Corbin (2001), Ho (2002)).

The majority of existing work on the savings-investment correlation concerns OECD countries with more limited work that addresses the case of LDCs. Moreover, the studies of OECD countries have generally found a large saving retention coefficient compared to a much weaker association between domestic savings and investment in the case of LDCs (Coakely *et al*. (1999), Ho (2002)). The purpose of this paper is to assess the extent of capital mobility for a large sample of LDCs over the study period of 1979-2001. It is possible that low test power is responsible for identifying a weak association between LDC domestic savings and investment. As argued below, the possibility of low test power justifies the employment of panel data cointegration tests. For LDCs, an investigation into the extent of capital mobility is of interest for a number of reasons. First, in the context of foreign capital-led economic growth, capital mobility means that a saving/investment constraint on economic growth can be lifted. Moreover, under the neoclassical growth model, a relatively poor economy with a low capital-labour ratio will be characterised by a relative high marginal product of capital and this can trigger greater savings and investment. Even if domestic savings are insufficient to respond to the high rates of return, the flow of foreign capital will ensure that the rate investment will remain high.[1] Of course, the converse case of capital flight can mean that investment is retarded.

Second, under the Mundell-Fleming (MF) model, perfect capital mobility means that fiscal (monetary) policy is most effective under a floating (fixed) exchange rate. From the point of view of policy-makers, the extent of capital mobility may play a role in the design of macroeconomic policy.

Third, there are issues of general macroeconomic instability associated with capital mobility. There have at times been massive increases in the flows of capital to LDCs (see, *inter alia*, Dooley *et al*. (1996)). Much of this capital is short-term in nature and brings with it many problems. For example, increased capital inflows can cause a real appreciation of the domestic currency and negate the impact of a nominal devaluation. The receiving country's currency may be subject to a speculative attack if the short-term capital inflows are volatile. When the capital account drives the current account deficit, large capital inflows can lead to an unsustainable current account deficit, particularly if the capital inflow is in response to a consumption boom. Capital flight can also have a destabilising effect on interest rates, exchange rates, foreign exchange reserves and money demand (as a result of currency substitution being linked with capital flight).

Given the importance of measuring LDCs capital mobility and the lack of existing research in this area, this study applies panel data cointegration techniques against a background of limited data. A limited time series sample size reduces the power of cointegration tests thereby making it harder to reject the null hypothesis of non-cointegration between domestic savings and investment (perfect capital mobility) against cointegration alternatives. Creating a larger panel data set from the individual time series increases the power of test and lessens the likelihood that cointegration has been rejected because the sample size is small. As well as using more observations, the panel data cointegrating techniques exploit the cross-

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[1] In the spirit of this logic, the basic Harrod-Domar growth model can be extended to the open economy.

country variations of the data in estimation. This is the first study to address capital mobility in the case of LDCs through the application of panel data cointegration techniques. Not only can we test the null of non-cointegration between domestic saving and investment (perfect capital mobility) against the null of cointegration (capital mobility), we are also able to formally test the null hypothesis of a long-run unity restriction on the saving retention coefficient that implies perfect capital immobility.

The paper is structured as follows. The following section discusses the methodological issues that are relevant to this study. We consider the interpretation of the basic F-H equation and outline the procedure used for the panel cointegration tests and fully modified ordinary least squares (FMOLS) estimation. The particular panel data cointegration methodology that we employ is based on Pedroni (1999, 2001, 2004). The third section discusses the data and results. We find that LDC capital is imperfectly mobile and that the degree of capital mobility is the same for both Asian and Latin American countries. The final section concludes.

II. Methodological Issues

The standard F-H equation may be written as follows:

$$I_t = \alpha + \beta S_t + \mu_t \tag{1}$$

where I is investment and S is domestic savings (both expressed as a percentage of GDP). The saving retention coefficient is the proportion of incremental savings that is invested domestically and is denoted by β where β=0 (β=1) implies perfect capital mobility (immobility) whereas 0<β<1 implies imperfect capital mobility.

There are, however, several challenges to the F-H interpretation of international capital mobility. First, the persistent correlation between saving and investment may be not due so much to imperfect capital mobility than to the pro-cyclical character of saving and investment in a real business cycle model. While a number of authors have undertaken cross-sectional analysis on sample averages in the period thus eliminating the influence of these cycles in the saving–investment correlation, permanent investment must equal permanent saving plus some constant in the long run. If period averaging is long enough, there might be a misspecification because one is estimating an identity in which the saving retention coefficient is equal to unity. Second, proceeding to estimate the F-H equation through time-series approaches can be subject to simultaneous equations bias.

Third, a high correlation between savings and investment may actually be indicative of high capital mobility. For example, Murphy (1984) and Baxter and Crucini (1993) suggest that a high domestic saving–investment correlation reflects the country's financial size in the world economy. When the country's financial system is highly developed in international terms, exogenous variations in domestic saving and investment rates affect world interest rates and induce joint movements in domestic saving and investment rates. In this context, the F-H test can be interpreted as a joint test of the hypothesis of capital mobility and the size of a country's financial system. In the case of an LDC with a relatively underdeveloped financial system, however, this argument is likely to be of less relevance than is the case for developed economies. Fourth, a new interpretation of the high saving-investment correlation being suggested is through the acknowledgement of a solvency constraint (no Ponzi financing). In the long term, the intertemporal budget constraint (IBC) is an indicator of a country's solvency expressed in terms of current account constancy, which can be interpreted in the Feldstein-Horioka approach as evidence of imperfect capital mobility (Coakley *et al.*, 1996). Moreover, studies that identify a stationary current account balance reflect a saving retention ratio of 1. Thus a high correlation between saving and investment may be a reflection of a country satisfying its IBC rather than evidence of zero capital mobility.[2]

This study tests for cointegration between investment and domestic savings. A major obstacle to doing this with respect to time-series data is the lack of observations on S and I. As argued above, one way forward is to conduct the investigation on the basis of a panel data set. For the purpose of cointegration testing within the panel of LDCs, we follow Pedroni (1999) who proposes a range of statistics that can be used to determine the presence of cointegration in heterogeneous panels.[3] These tests do not constrain the estimated slope coefficients to be same across the panel and are applicable where regressors are fully

[2]Using data for LDCs over a 1956-90 sample period, Coakely *et al.* (1999) identify current stationarity in 14/44 cases thereby providing evidence that less than one-third of LDCs conform to the long-run solvency constraint. In addition to this, they confirm current account stationarity in 12/23 OECD economies.

[3]Corbin (2001) highlights the importance of controlling for the heterogeneity of countries in a cross-section analysis of the saving–investment correlation for a group of countries using panel data. It is argued that the individual and temporal dimensions of the data enables the estimation of the coefficient of the saving–investment correlation. Moreover, obtaining a high coefficient of correlation in the cross-section analysis, may be less due to the existence of a common characteristic affecting all the countries in the sample in the same way in a given period (imperfect capital mobility) than to the existence of specific individual country effects.

endogenous.[4] Within a panel setting, the test statistics are constructed using the residuals from the following hypothesized cointegrating regression based on equation (1),

$$I_{it} = \alpha_i + \beta_i S_{it} + \mu_{it} \tag{2}$$

where α_i allows the cointegrating regression to include country-specific fixed effects. The procedure for computing the test statistics involves estimating the hypothesized cointegration regression described in (2) and using the residuals μ_{it} to estimate the appropriate autoregression. From this, one may compute the panel ADF statistic which is a parametric statistic and analogous to the Levin and Lin (1993) panel data unit root test applied to the estimated residuals of cointegrating regression.[5] This statistic is referred to as a *within-dimension* statistic that effectively pools the autoregressive coefficients across different countries during the unit root test. A common value for the autoregressive coefficient is specified under the alternative hypothesis of cointegration. The second statistic is based on a group mean approach. The group ADF statistic is a parametric statistic and analogous to the Im *et al.* (1997, 2003) test for a unit root panel that is applied to the estimated residuals of a cointegrating regression. This statistic is referred to as a *between-dimension* statistic that averages the estimated autoregressive coefficients for each country. Under the alternative hypothesis of cointegration, the autoregressive coefficient is allowed to vary across countries. This allows one to model an additional source of potential heterogeneity across countries. Following an appropriate standardization, both of these statistics will be distributed as standard normal as both the number of observations (N) and number of LDCs in the panel (T) grow large. Both of these statistics diverge to negative infinity under the alternative hypothesis and consequently the left tail of the normal distribution is used to reject the null hypothesis of non-cointegration.

Having tested for cointegration, the second stage of the investigation is to analyse equation (2). Pedroni (2001, 2004) describes how FMOLS procedures can be employed to obtain the panel data estimates for β_i. Using a dynamic modelling

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[4]See Pedroni (2004).

[5]In the case of the panel ADF statistic, (2) is also run in first difference form where the estimated residuals are saved. The long–run variance of is computed and used as a nuisance parameter estimator in the computation of the test statistic.

procedure results in a more powerful test for cointegration as well as giving generally unbiased estimates of the long-run relationship and standard *t*-statistics. FMOLS amounts to the application of non-parametric adjustment to the OLS estimates of both the long-run parameter β_i and associated *t*-statistic, on account of any bias due to autocorrelation or endogeneity bias that shows up in the OLS residuals [Phillips and Hansen (1990)].

Following on from (2), let $\xi_{it} = (\hat{\mu}_{it}, \Delta S_{it})'$ be a stationary vector comprising the estimated residuals and the differences in savings. Also, let $\Omega_i = lim_{T\to\infty} E\left[T^{-1}\left(\sum_{t=1}^{T}\xi_{iT}\right)\left(\sum_{t=1}^{T}\xi_{iT}\right)'\right]$ be the long-run covariance for this vector process which can be decomposed into $\Omega_i = \Omega_i^0 + \Gamma_i + \Gamma_i'$ where Ω_i^0 is the contemporaneous covariance and Γ_i is a weighted sum of autocovariances. Pedroni shows that the group mean panel FMOLS estimator is given as

$$\hat{\beta}^*_{GFM} = N^{-1}\sum_{i=1}^{N}\left(\sum_{t=1}^{T}(S_{it}-\bar{S}_i)^2\right)^{-1}\left(\sum_{t=1}^{T}(S_{it}-\bar{S}_i)I_{it}^* - T\hat{\gamma}_i\right) \tag{3}$$

where $I_{it}^* = (I_{it} - \bar{I}_i) - \frac{\hat{\Omega}_{21i}}{\hat{\Omega}}\Delta S_{it}$ and $\hat{\gamma}_i \equiv \hat{\Gamma}_{21i} + \hat{\Omega}^0_{21i} - \frac{\hat{\Omega}_{21i}}{\hat{\Omega}_{22i}}(\hat{\Gamma}_{22i} + \hat{\Omega}^0_{22i})$. The between-dimension estimator is calculated as $\hat{\beta}^*_{GFM} = N^{-1}\sum_{i=1}^{N}\hat{\beta}^*_{FM,i}$ where $\hat{\beta}^*_{FM,i}$ is the conventional FMOLS estimator applied to the i^{th} member of the panel. The associated *t*-statistics are calculated as

$$t_{\hat{\beta}^*_{GFM}} = N^{-0.5}\sum_{i=1}^{N} t_{\hat{\beta}^*_{FM,i}} \tag{4}$$

where. $t_{\hat{\beta}^*_{FM,i}} = (\hat{\beta}^*_{FM,i} - \beta_0)\left(\hat{\Omega}^{-1}_{11i}\sum_{t=1}^{T}(S_{it}-\bar{S}_i)^2\right)^{0.5}$.

In the empirical analysis, we focus on the between-dimension panel FMOLS tests. There are several advantages over the within-dimension approach. First, the between-dimension approach allows for greater flexibility in the presence of heterogeneity across the cointegrating vectors where β_i is allowed to vary. Under the within-dimension approach, β_i would be constrained to be the same value for each country under the alternative hypothesis. Second, the point estimates of the between dimension estimator can be interpreted as the mean value of the cointegrating vectors. This is helpful in the interpreting the results. Third, the between-dimension estimator suffers from lower small-sample size distortions than is the case with the within-dimension estimator.

III. Data and Results

The study initially employs annual data on investment and domestic savings (expressed as a percentage of GDP) for twenty-four LDCs over the study period 1979-2001 inclusive. The study period is dictated by data availability across this large sample that comprises Barbados, Brazil, Chile, Colombia, Costa Rica, Ecuador, Egypt, El Salvador, Guatemala, Honduras, India, Jamaica, Kenya, Mexico, Morocco, Nigeria, Pakistan, Philippines, Singapore, South Africa, Sri Lanka, Thailand, Uruguay and Venezuela. All data are taken from the *World Development Indicators* via the World Bank database. First of all, the stationarity of these series is investigated using univariate ADF unit root testing. Table 1 reports that non-stationarity in the investment and/or savings rates is rejected at the 5 per cent significance level or stronger in eight cases namely, Brazil, Ecuador, Guatemala, India, Kenya, Nigeria, Uruguay and Venezuela.[6] These countries are excluded from the main cointegration analysis that should be based on long-run equilibrium relationships between non-stationary series. This still leaves more than half the sample characterized by non-stationary investment and saving rates namely, Barbados, Chile, Colombia, Costa Rica, Egypt, El Salvador, Honduras, Jamaica, Mexico, Morocco, Pakistan, Philippines, Singapore, South Africa, Sri Lanka and Thailand. These sixteen countries are therefore included in the panel data cointegration analysis. Table 2 reports the Pedroni cointegration tests based on equation (2). These tests include time-specific dummies to allow for the possibility that residuals are correlated across countries. The null of non-cointegration is strongly rejected at the 1 per cent significance level by both the panel ADF and group ADF tests respectively. The former test is relatively more restrictive in that a common value for the autoregressive coefficient is specified under the alternative hypothesis of cointegration. The group ADF test allows for the autoregressive coefficient to vary across countries under the alternative hypothesis.

Table 3 reports the FMOLS panel estimates of the cointegrating relationships between investment and domestic saving. For the individual country estimates,

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[6]The lag lengths in these ADF regressions were determined by the AIC. There was no affect on the qualitative conclusions drawn in this study if alternative methods for lag selection such as the BIC were employed.

Table 1. ADF Unit Root Tests for Individual Countries

	S		*I*	
	ADF	ADF (trend)	ADF	ADF (trend)
Barbados	-2.662	-3.434*	-1.818	-1.507
Brazil	-1.824	-1.809	-4.081***	-4.283***
Chile	-2.083	-3.447	-1.412	-1.969
Colombia	-1.648	-1.788	-2.863	-2.920
Costa Rica	-1.448	-1.690	-2.879*	-3.168
Ecuador	4.939	4.612	-3.003**	-2.503
Egypt	-2.519	-2.436	-1.592	-2.872
El Salvador	-2.416	-2.840	-1.905	-2.943
Guatemala	-3.324**	-3.072	-2.203	-2.765
Honduras	-0.868	-2.880	-1.198	-2.966
India	-2.474	-4.560***	-2.223	-1.726
Jamaica	-1.349	-2.083	-1.822	-3.123
Kenya	-1.715	-2.335	-3.483**	-3.552*
Mexico	-2.008	-2.283	-2.874*	-2.760
Morocco	-1.488	-2.218	-2.371	-2.282
Nigeria	-3.185**	-3.244*	-2.862*	-2.802
Pakistan	-2.308	-2.678	0.631	-1.090
Philippines	-2.155	-1.888	-2.344	-2.495
Singapore	-1.952	-2.353	-2.317	-3.374*
South Africa	-1.381	-2.628	-1.245	-1.256
Sri Lanka	-2.868*	-3.331*	-2.831*	-3.026
Thailand	-1.924	-1.525	-1.263	-0.682
Uruguay	-3.002**	-3.399*	-7.880***	-7.223***
Venezuela	-3.332**	-3.430*	-3.191**	-3.730**

Notes for Table 1. These tests employ annual data on domestic investment and savings expressed as a percentage of GDP for the study period 1979-2001. In all cases, the lag length is selected according to the AIC. ***, ** and * respectively denote rejection of the non-stationary null at the 1, 5 and 10 per cent significance level respectively. The 1, 5 and 10 per cent critical values are -3.75, -3.00 and -2.63 respectively for the nontrended ADF tests and -4.38, -3.60 and -3.24 respectively in the trended ADF tests.

Table 2. Panel Data Cointegration Tests

Test	Test Statistic
panel ADF	-3.341***
group ADF	-3.477***

Notes for Table 2. These are the Pedroni tests for panel non-cointegration [discussed in Pedroni (1999)] between investment and saving rates. The panel ADF statistic is a within-dimension statistics, the group ADF statistic is a between-dimension statistic. Both tests are asymptotically normal. *** denotes rejection of the non-cointegration null at the 1 per cent significance level with a critical value of –2.326.

Table 3 reports $\beta_i > 0$ in thirteen cases but there exists considerable variation in the country-by-country experiences. The null of a zero slope coefficient is rejected at

Table 3. FMOLS Estimation of the Cointegrated Panel

	$\hat{\beta}^*_{FM,i}$	$t_{\hat{\beta}^*_{FM,i}}$ $H_0: \beta^*_{FM,i} = 0$	$t_{\hat{\beta}^*_{FM,i}}$ $H_0: \beta^*_{FM,i} = 1$
Barbados	0.170	0.461	-2.249
Chile	0.520	8.844	-8.153
Colombia	0.004	0.016	-4.119
Costa Rica	0.010	0.080	-7.612
Egypt	-0.205	-0.521	-3.061
El Salvador	0.623	5.637	-3.407
Honduras	0.505	4.938	-4.843
Jamaica	0.698	6.245	-2.702
Mexico	0.318	1.258	-2.695
Morocco	-0.131	-0.623	-5.381
Pakistan	0.106	0.737	-6.197
Philippines	0.660	1.832	-0.945
Singapore	-0.382	-1.211	-4.385
South Africa	1.240	4.880	0.944
Sri Lanka	0.194	0.587	-2.435
Thailand	1.122	2.822	0.307
	$\hat{\beta}^*_{GFM}$	$t_{\hat{\beta}^*_{GFM}}$ $H_0: \beta^*_{GFM} = 0$	$t_{\hat{\beta}^*_{GFM}}$ $H_0: \beta^*_{GFM} = 1$
Group	0.341	8.995	-14.233

Notes for Table 3. This table reports FMOLS panel data estimates of βi in equation (2) using the Pedroni panel data cointegration methodology. These estimates include common time dummies. The bottom row refers to the group-mean estimates. Each estimate of β is accompanied by two t-statistics. Column 3 reports t-statistics for the null β=0, while column 4 reports t-statistics for the null β=1. All t-statistics are asymptotically normal. The t-statistics are asymptotically normal with 1, 5 and 10 per cent critical values of +/-2.57, +/-1.96 and +/-1.64 respectively.

the 5 per cent significance level in all cases except Barbados, Colombia, Costa Rica, Egypt, Mexico, Morocco, Pakistan, Philippines, Singapore and Sri Lanka. At the 5 per cent significance level, the null of a unity slope coefficient and therefore perfect capital immobility is rejected in the all cases except Philippines, South Africa and Thailand. According to the group mean estimates, a long-run relationship between S and I is confirmed with β=0.341. At the 5 per cent significance level, this coefficient is both significantly different from zero and unity with t-statistics of 8.995 and –14.233 respectively (see columns 2 and 3). This result indicates that capital mobility is imperfect where with a savings retention ratio that is closer to zero than unity. In comparing this results with existing panel cointegration studies of capital mobility, one may reflect on the study by Ho (2002) who uses panel cointegration techniques to estimate the F-H equation in the case of twelve OECD economies.[7] Using data for 1950-92, Ho produces a panel estimate of β=0.786 suggesting that capital is actually less mobile for the OECD economies than for LDCs over the 1979-2001 period. Also, one might note the study by Coakely *et al*. (1999) who employ data for forty-four LDCs over a study period that runs upto 1990. While panel data cointegration techniques are not employed, they find that the average time-series estimate for β_1 based on the Johansen maximum likelihood procedure is 0.44 while the cross-sectional estimate of β is 0.40. In the case of the cross section of OECD economies, the cross-sectional estimate of β is 0.732 also suggesting that capital mobility is greater for LDCs.

In this study, the null hypothesis of β=1 is rejected for the group mean estimate thereby rejecting long-run current account sustainability through implication.[8] A non-stationary current account is inconsistent with the sustainability of external debts and might indicate there is an incentive for the country to default on its external debts. Furthermore, the non-stationarity of the current account is inconsistent with the intertemporal model of the current account, and hence refutes its validity.[9] The modern intertemporal model of current account determination uses consumption-smoothing behavior to predict that the current account acts as a buffer to smooth consumption in the face of shocks. This implies that the current account should in fact be a stationary rather than a non-stationary series.

A further interesting issue is whether there exists evidence that capital mobility differs significantly across continents. From the full panel of sixteen LDCs, Pakistan, Philippines, Singapore, Sri Lanka and Thailand are used to form an Asian panel, while Chile, Colombia, Costa Rica, El Salvador, Honduras and Mexico are used to form a Latin American panel. Tables 4 and 5 present panel cointegration tests and FMOLS group estimates of β for these sub-groups. Both the panel ADF and group ADF statistics reported in Table 4 indicate that the null of non-cointegration is strongly rejected for both regional groupings. Table 5 reports

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[7]In this study, Ho employs the earlier panel data cointegration test advocated by Kao (1999) which is a residual-based test that imposes a common slope coefficient. By contrast, the Pedroni *group ADF* test allows the estimated slope parameters to vary across individual members of the panel (Pedroni (2004)).

[8]Compared to existing work on OECD countries, the examination of the sustainability of LDC current account deficits is a relatively unexplored area. Other recent studies of LDCs include *inter alia* Pattichis and Kanaan (2001), who find that the trade balance is stationary and therefore sustainability in the case of Lebanon, and Coakley and Kulasi (1997) who confirm likewise for India but not Korea and Taiwan.

[9]See, for example, Husted (1992) and references therein.

Table 4. Panel Data Cointegration Tests for Sub-groups

Test	Asia	Latin America
panel ADF	-2.260**	-3.402***
group ADF	-2.504***	-3.400***

Notes for Table 4. These are the Pedroni tests for cointegration [discussed in Pedroni (1999)] between investment and saving rates. The panel ADF statistic is a within-dimension statistics, the group ADF statistic is a between-dimension statistic. Both tests are asymptotically normal. *** and ** denote rejection of the non-cointegration null at the 1 per cent and 5 per cent significance levels with respective critical values of -2.326 and -1.64.

Table 5. FMOLS Estimation of the Cointegrated Panel for Sub-groups

	$\hat{\beta}^*_{GFM}$	$t_{\hat{\beta}^*_{FM,i}}$ $H_0: \beta^*_{GFM} = 0$	$t_{\hat{\beta}^*_{FM,i}}$ $H_0: \beta^*_{FM,i} = 1$	$t_{\hat{\beta}^*_{FM,i}}$ $H_0: \beta^*_{FM,i} = 0.341$
Asia	0.340	2.132	-6.107	-0.678
Latin America	0.330	8.480	-12.586	1.297

Notes for Table 5. This table reports FMOLS panel data estimates of β_1 in equation (2) using the Pedroni panel data cointegration methodology. These estimates include common time dummies. The bottom row refers to the group-mean estimates. The estimate of β is accompanied by three t-statistics. Column 3 reports t-statistics for the null β=0, column 4 reports t-statistics for the null β=1 and column 5 reports t-statistics for the null β=0.341. The t-statistics are asymptotically normal with 1, 5 and 10 per cent critical values of +/-2.57, +/-1.96 and +/-1.64 respectively.

FMOLS group estimates of β=0.340 and β=0.330 for the Asian and Latin American groups respectively. In both cases, the respective nulls of zero and unity are rejected confirming imperfect capital mobility for these LDCs but the estimated saving retention ratios are very slighter lower than for the full panel. This suggests that capital mobility may be slightly greater for the Asian and Latin American LDCs, however Table 5 also indicates that the null β=0.341, which is based on the saving retention ration across the entire sample of LDCs, is easily accepted at the 5 per cent significance level for both sub-groups.

IV. Summary and Conclusion

This is the first study of long-run capital mobility with respect to LDCs using panel data cointegration methods. The assessment is based on estimating the long-run relationship between investment and savings as originally proposed by Feldstein and Horioka and others in the measurement of OECD capital mobility. Using annual data for a panel of sixteen LDCs over the study period 1979-2001, a long-run cointegrating relationship between domestic savings and investment is strongly confirmed and fully modified ordinary least squares estimation indicates that capital is imperfectly mobile. The low saving retention coefficient of about one-third suggests that capital mobility may actually be fairly high for LDCs. Further panel estimation for sub-groups comprising Asian and Latin American countries offer little evidence that the extent of capital mobility is significantly different between these panels.

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