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Regional integration, capital mobility and financial intermediation revisited: Application of general to specific method in panel data

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ABSTRACT

We utilize the Feldstein–Horioka puzzle to investigate the impact of regional integration agreements (AFTA, EU, EFTA, CARTAGENA, MERCOSUR and NAFTA) on the international capital mobility. In doing so, we employed a novel empirical technique i.e. the general to specific (GETS) method of Hendry (1995) to estimate the cointegrating equation and dynamic adjustments in panel data. Using the classical fixed and random effects estimators, we estimate the long- and short-run effects in the same model and we show that it is possible to estimate the lagged adjustment process. The procedure used is general enough to allow for the presence of endogeneity, heteroscedasticity, serial correlation and cross-sectional dependence in the residuals. Our findings show that the estimate of saving retention has declined and the speed of adjustment has increased in the post-integration period, implying that the international mobility of capital has increased in these countries. Moreover, our findings reveal that regional integrations stimulate financial intermediation, which in turn, improves real productivity.

1. Introduction

The relationship between investment and saving has been extensively analyzed since the publication of the seminal article of Feldstein and Horioka (1980) (henceforth *FH*). According to *FH*, high correlation between domestic savings and investment implies that the mobility of capital is low; this relationship is well known as the Feldstein–Horioka puzzle. Using the cross-section data of 16 OECD countries over the period 1960–1974, they estimated the relationship between domestic investment share of GDP (*ITY*) and domestic saving share of GDP (*STY*). The null hypothesis is that under complete capital mobility the savings retention coefficient should be zero. They found that this estimate is close to unity and this implies that domestic saving is the main source of funds for investment.

Testing the *FH* puzzle using cross-section data was extremely popular in the 1980s and 1990s and majority of the studies appeared to confirm the original *FH* result and revealed that the savings retention estimate for the OECD countries did not decline when data was extended up to 1980 and beyond, for instance among others were Feldstein (1983), Murphy (1984), Penati and Dooley (1984), Obstfeld (1986, 1995), Dooley et al. (1987), Golub (1990), Tesar (1991), Artis and Bayoumi (1991),

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Sinn (1992) and Coakley et al. (1996). A few studies found that the savings retention coefficient had only marginally declined (Feldstein and Bacchetta, 1991; Coakley et al., 1996; Artis and Bayoumi, 1991).

Since the 1990s, the application of non-stationary time series and panel methods has become of special interest in testing the validity of the *FH* puzzle. The first group of studies have followed the time series estimation route through using the country-specific time series techniques. Among them, we can highlight contributions by Miller (1998), Alexakis and Apergis (1992), Argimon and Roldan (1994), Ghosh (1995), Goldberg et al. (1995), Coakley et al. (1996), Jansen (1996), Liu and Tanner (1996), Hussein (1998), Kim (2001), Schmidt (2003), Hoffmann (2004) and Kumar et al. (2012). These studies provide evidence of non-stationarity in *ITY* and *STY* i.e. integrated of order one $I(1)$ and therefore they support the use of nonstationary cointegration techniques to investigate the *FH* puzzle.[1] Obstfeld (1986) and Coakley et al. (1998) suggest that the country-specific time series estimates are vital especially as a guide for pooling time series observations from different countries.

The second group of studies investigated the *FH* puzzle using panel methodologies (Coakley et al., 1999, 2001, 2004; Krol, 1996; Coakley and Kulasi, 1997; Oh et al., 1999; Coiteux and Olivier, 2000; Cadoret, 2001; Corbin, 2001; Ho and Chiu, 2001; Kim, 2001; Tsung-wu, 2002; Pelgrin and Schich, 2004; Kim et al., 2005; Bahmani-Oskooee and Chakrabarti, 2005; Chakrabarti, 2006; Murthy, 2007; Christopoulos, 2007; Di Iorio and Fachin, 2007; Herwartz and Xu, 2009; Fouquau et al., 2009). The results in these studies reinforce the previous results based on a time series framework. They attained lower saving retention estimates, although the relationship between saving and investment is shown to be stronger for OECD than the less developed countries. In spite of varied panel techniques applied, these studies all start from the same premise that the saving and investment are non-stationary processes.

While many accept the high association between investment and saving, controversy remains with the interpretation of the savings retention coefficient; in particular how informative is this estimate about capital mobility. Ghosh (1988), Obstfeld (1986), Uctum and Wickens, 1990, Genberg and Swoboda (1992) and Argimon and Roldan (1994) suggest that capital mobility should be examined within the inter-temporal model of saving and investment behaviour. Baxter and Crucini (1993) argued that in the long run, technological variables and the demographic structure of the population could drive investment and saving, thereby inducing positive correlation even with perfect capital mobility. According to Jansen (1996, 1998), Coakley and Kulasi (1997) and Pelgrin and Schich (2004), the long run relationship between the investment and saving could be treated as solvency conditions that must be satisfied. Byrne et al. (2009) have argued that this correlation could be due to common global shocks and therefore it may not be interpreted as evidence against capital mobility. Nevertheless, we take the view of many that the *FH* puzzle is a simple and an indirect test on the extent to which capital is mobile across countries. If tested for structural breaks (for example, considering regional economic integration in the sample) it may also give an indication about changes in capital mobility. Obstfeld and Rogoff (2000) have called *FH* puzzle the mother of all puzzles because it provides useful insights on the international capital mobility.

In our view, the empirical *FH* puzzle literature is notoriously fragile because in most studies the unit root and stationarity hypothesis testing ignored to consider the presence of structural breaks. Perron (1989, 1997) showed that the ability to reject a unit root decreases when the stationary alternative is true and an existing structural break is ignored. Carrion-i-Silvestre et al. (2005) pointed out that this kind of misspecification error can lead to spurious non-stationarity. However, there exist a few studies that have employed the structural break tests to determine breaks in the cointegrating relationship of investment and saving. Westerlund (2006) considers the presence of multiple breaks in a sample of 15 OECD countries and found that saving and investment are cointegrated under the presence of level and trend shifts. Using the structural break test in Westerlund (2006), Kumar and Rao (2011) found that structural changes did reduce the savings retention estimates in OECD countries, especially in the post Bretton Woods and Maastricht periods. Analogous inferences were made by Rao et al. (2010) using the exogenous structural break tests in Mancini-Griffoli and Pauwels, 2006. Di Iorio and Fachin (2007) have used panel bootstrap tests to examine the *FH* puzzle for a sample of 12 EU countries; their results show that the bootstrap panel stability tests allow for cointegration between saving and investment in the long run with at least one break.

This paper investigates the stationarity properties of investment and saving series for panels that comprise six regional investment agreements.[2] The contribution of this paper is threefold. Firstly, we utilize the panel stationarity test of Carrion-i-Silvestre et al. (2005) to test for integrated order of the variables. The innovative aspect of this test is that it allows for cross-sectional dependence (CSD) and multiple structural breaks. Our results suggest that *ITY* and *STY* series can be characterized as stationary processes evolving around a broken trend. Our results reveal that regional economic integrations are a major source of structural change in the *ITY* and *STY* series. This result casts doubt on almost all recent empirical studies on *FH* that utilized nonstationary time series methods to test the puzzle. Secondly, we provide empirical evidence on how international capital mobility has been affected by various regional investment agreements. In doing so, we employ the London School of Economics (LSE) Hendry's (1995) general to specific (GETS) method to estimate the savings retention coefficients in a panel framework. The procedure used is general enough to allow for the presence of endogeneity, heteroscedasticity, serial correlation and CSD in the residuals. GETS specifications have received limited attention in panel data estimations; therefore in this paper we use GETS approach to estimate the short run dynamic adjustment equations with panel data. Since the *ITY*

[1] Some studies did use rates of investment and savings instead of the ratios; for a comprehensive survey see Apergis and Tsoumas (2009).

[2] ASEAN Free Trade Area (AFTA), European Union (EU), European Free Trade Association (EFTA), Codification of the Andean Sub-regional Integration Agreement (CARTAGENA), Southern Common Markets (MERCOSUR) and North American Free Trade Agreement (NAFTA).

Table 1
Regional investment integrations, year in force and countries.

Regional investment agreements			
Region	Agreement	Year in force	Countries
ASIA	AFTA	1992	Brunei,[a] Indonesia, Malaysia, Philippines, Singapore, Thailand, Vietnam,[a] Laos,[a] Myanmar[a] and Cambodia.[a]
EUROPE	EU	1993	Austria, Belgium, Bulgaria,[a] Cyprus, the Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Latvia, Lithuania,[a] Luxembourg, Malta, the Netherlands, Poland, Portugal, Romania, Slovakia, Slovenia,[a] Spain, Sweden and the UK.
	EFTA	1960	Liechtenstein, Iceland, Norway and Switzerland.
AMERICA	CARTAGENA	1969	Bolivia, Columbia, Ecuador, Peru and Venezuela.
	MERCOSUR	1991	Argentina, Brazil, Paraguay and Uruguay.
	NAFTA	1994	Canada, Mexico and the USA.

[a] The country is not included in our sample due to unavailability of data or the country has joined the agreement much later. Vietnam, Laos, Myanmar and Cambodia joined AFTA much later.

and *STY* series are stationary, it is therefore pragmatic to apply the GETS technique; a nonstationary panel cointegration method (for example, Mark and Sul, 2003; Pedroni, 2004; Westerlund, 2007) may be unfeasible. Lastly, we investigate the relationships between regional integration, financial intermediaries and productivity. Much of the empirical literature has focused on either the *FH* puzzle and international capital mobility or financial development and economic growth. This paper makes the first attempt in this literature to explore the associations between integration, financial intermediaries and productivity for a wide range of countries.

The remainder of the paper is structured as follows. Section 2 provides a discussion of the data. Section 3 details the methods used to analyze the *FH* puzzle. Section 4 presents the empirical results. Finally, Section 5 concludes.

2. Data

Our sample consists of 44 countries (see Table 1) for which annual data on *ITY* (gross domestic investment as a share of GDP) and *STY* (gross domestic savings as a share of GDP) are available from 1960 to 2012. The averages of *ITY* and *STY* for the panel are 30.03% and 26.39%, respectively. At the same time the standard deviations for the two variables are 12.24% and 14.95%, respectively. For both variables, the standard errors are low at around 0.2.

Fig. A.1 in Appendix illustrates the averages of *ITY* and *STY* for each country operating under various agreements over the period 1960–2012. It is observed that most countries have investment and savings ratio averages more than 20%, except for Bolivia and Ecuador for CARTAGENA, USA for NAFTA, Paraguay (only *STY*) and Uruguay for MERCOSUR, Cyprus (only *STY*), Greece (only *STY*), Portugal (only *STY*), Romania (only *STY*) and the UK for EU. Fig. A.2 in Appendix depicts the *ITY* and *STY* averages for each panel corresponding to the agreements.

Other series that have been used in this paper are narrow money as a percentage of GDP (*M1Y*), broad money as a percentage of GDP (*M2Y*), liquid liabilities as a percentage of GDP (*LLY*), domestic credit to private sector by banks as a percentage of GDP (*DCPSY*) and real productivity (*y* is real GDP divided by the total number of workers). All data is attained from the World Development Indicators (WDI) and International Financial Statistics (IFS).

3. Methodology

3.1. Panel unit root test without structural breaks

We first utilize a panel unit root test without structural breaks i.e. Hadri (2000). This is a residual based Lagrange multiplier (LM) first generation panel unit root test. The null hypothesis of stationarity in all panel units is tested against the alternative hypothesis of a unit root in all cross-section units. In order to test the null hypothesis of stationarity, Hadri (2000) proposes to use the panel version of the test in Kwiatkowski et al. (1992) applied in the univariate context. In its heterogeneous version the test statistic is given by:

$$\eta_k = N^{-1}\sum_{i=1}^{N}\left(\hat{\omega}_i^{-2}T^{-2}\sum_{t=1}^{T}S_{i,t}^2\right) \tag{1}$$

where $k = \{\mu, \tau\}$. $S_{i,t} = \sum_{j=1}^{t}\hat{\varepsilon}_{i,j}$ denotes the partial sum process obtained from the estimated OLS residuals when regressing the individual time series on a constant or on a time trend. $\hat{\omega}_i^2$ is the consistent estimate of the long run variance of $\varepsilon_{i,t}$. The test statistic is distributed as standard normal under the null hypothesis. The series may be stationary around a deterministic level (specific to the unit – i.e. a fixed effect) or around a unit-specific deterministic trend. The error process may be assumed to be homoscedastic across the panel, or heteroscedastic across units. Serial dependence in the disturbances may also be taken into account using a Newey–West estimator of the long-run variance. The residual-based test is based on the squared partial sum process of residuals from demeaning (detrending) model of level (trend) stationarity.

3.2. Panel unit root test with structural breaks

We apply Carrion-i-Silvestre et al. (2005) test to analyze stationarity in *ITY* and *STY* series. This test allows for the presence of multiple structural breaks and CSD in the errors. Carrion-i-Silvestre et al. (2005) modified the Hadri (2000) test to allow for multiple structural breaks by including dummy variables in the deterministic specification of the model. Under the null hypothesis the data generating process for the variable is assumed to be (Carrion-i-Silvestre et al., 2005, p. 844):

$$y_{i,t}=\alpha_i+\sum_{k=1}^{mi}\theta_{i,k}DU_{i,k,t}+\beta_i t+\sum_{k=1}^{mi}\gamma_{i,k}DT^*_{i,k,t}+\varepsilon_{i,t} \tag{2}$$

where the dummy variable $DT^*_{i,k,t}=t-T^i_{b,k}$ for $t>T^i_{b,k}$ and 0 elsewhere, $k=\{1,\ldots,m_i\}$, $m_i\geq 1$. The model in (2) includes individual effects and individual structural break effects, i.e. shifts in the mean caused by the structural breaks. It also includes temporal effects (if $\beta_i\neq 0$) and temporal structural break effects (if $\gamma_{i,k}\neq 0$), i.e. when there are shifts in the individual time trend. According to Carrion-i-Silvestre et al. (2005, p. 844), this model is general enough to allow for the structural breaks to have different effects on each individual time series and to be located at different dates. In addition, it permits the individuals to have a different number of structural breaks. Following Hadri (2000), Carrion-i-Silvestre et al. (2005) formulate the test of the null hypothesis of a stationary panel as:

$$\mathrm{LM}(\lambda)=N^{-1}\sum_{i=1}^{N}\left(\hat{\omega}_i^{-2}T^{-2}\sum_{t=1}^{T}S^2_{i,t}\right) \tag{3}$$

where $S_{i,t}=\sum_{j=1}^{t}\hat{\varepsilon}_{i,j}$ denotes the partial sum process attained from the estimated OLS residuals from Eq. (2). $\hat{\omega}_i^2$ is the consistent estimate of the long run variance of $\varepsilon_{i,t}$. λ in Eq. (3) denotes the dependence of the test on the dates of the break. Carrion-i-Silvestre et al. (2005) utilize the procedure in Bai and Perron (1998) to estimate the number of structural breaks and their positions.

3.3. GETS procedure in panel data

The LSE Hendry's GETS approach is based on a dynamic specification. GETS was developed because the econometricians at the LSE were concerned with the methodological conflict between the static nature of equilibrium relationships and the data that is used to estimate them. Economic theory seldom gives much information about dynamic adjustments and how long is the transition process in the real calendar time. However, the data from the real world which is used to test theories are hardly generated by an equilibrium world. Therefore, there is a methodological problem with using data generated from a disequilibrium world to test equilibrium theories. And this is the starting point for the development of GETS. Rao (2007) argued that GETS is appropriate to determine the dynamic adjustment structure by using the data itself so that these are consistent with the data generating process.

The first step in GETS procedure is to specify the general (unrestricted) autoregressive distributed lag (ARDL) model. In the case of *ITY* and *STY* relationship, the general GETS specification is as follows:

$$\begin{aligned}\Delta ITY_{it}&=\lambda[ITY_{it-1}-(\alpha+\beta STY_{it-1})]+\sum_{i=0}^{m}\gamma_i\Delta STY_{it-i}+\sum_{i=1}^{n}\Phi_i\Delta ITY_{it-i}+u_{it}\\ \Delta ITY_{it}&=\lambda ECM_{it-1}+\sum_{i=0}^{p}\delta_i\Delta STY_{it-i}+\sum_{i=1}^{n}\theta_i\Delta ITY_{it-i}+u_{it}\end{aligned} \tag{4}$$

where i and t are country and time indices, respectively, and $u_{it}\sim N(0,\sigma)$ for all i and t. Selecting the appropriate lag length is a problem and it is therefore useful to start with a large lag structure; it also depends on the time dimensions available in the data.[3] To this end, it is desirable to have sufficient T. The ECM denoted by λ consists of the levels of the variables (in square bracket) and it is $I(0)$. Hendry (1995) stated that if economic theory is correct then the combination of variables in the ECM should be stationary. In addition, the changes in the variables are also $I(0)$.

In the second step, Eq. (4) is estimated using the classical fixed effects (FE) and random effects (RE) estimators. The third step is to apply the standard variable deletion tests to attain the parsimonious GETS FE and RE versions. However, the long

[3] *PcGets* software uses multiple search paths to derive, often in a few seconds, a parsimonious congruent specification from a large and comprehensive general unrestricted model. *PcGets* contains an automatic panel data selection, taking account of outliers. *PcGive* software can also be used to achieve parsimonious model, however the panel options are not yet operational.

Table 2
Hadri panel unit root test 1960–2012.

Series	AFTA	EU	EFTA	CARTAGENA	MERCOSUR	NAFTA
ITY	15.820	1.534	7.002	11.871	8.900	10.025
	(0.00)***	(0.36)	(0.00)***	(0.00)***	(0.00)***	(0.00)***
STY	19.023	28.002	9.201	3.939	10.024	11.820
	(0.00)***	(0.00)***	(0.00)***	(0.13)	(0.00)***	(0.00)***
Δ*ITY*	1.159	1.560	0.663	1.983	1.160	0.012
	(0.40)	(0.35)	(0.61)	(0.28)	(0.39)	(0.99)
Δ*STY*	2.044	0.021	2.830	1.266	0.036	3.750
	(0.22)	(0.99)	(0.26)	(0.42)	(0.99)	(0.18)

NB: probability values are reported in the parentheses.
*** Significance level at 1% level.

run (error correction) variables are of main interest and therefore they are not deleted at any stage. The fourth step is to select between the FE and RE model. To this end, it may be important to perform the Hausman test. The null hypothesis is that the preferred model is RE against the alternative of FE, see Greene (2008). It tests whether the unique errors (u_i) are correlated with the independent variables. The null is that errors are not correlated with the independent variables versus the alternative that errors are correlated.

The fifth step attempts to address endogeneity issue in the model. We suggest a simple procedure that utilizes insights from the exogeneity tests of block Granger non-causality. Endogeneity is tested by regressing the first difference of each explanatory variable on the one period lagged ECM normalized on the dependent variable. There exists endogeneity problem if the one period lagged ECM is statistically significant with correct negative sign in the equation for first difference explanatory variable. To minimize endogeneity bias, instrumental variable estimators should be employed.

The final steps involve addressing the CSD in the errors and applying other diagnostic tests. If using the FE model, the Breusch–Pagan LM test may be used to identify the issue. To address CSD, we suggest use of the procedure in Driscoll and Kraay (1998). Their method produces standard errors for estimates estimated by the FE (within) regression. The error structure is assumed to be heteroscedastic, autocorrelated up to some lag and perhaps correlated between the groups (panels). The standard errors are robust to CSD especially when the time dimension becomes large. According to Hoechle (2007) this procedure offers no restrictions on the limiting behaviour of the number of panels. Further, the size of the cross-sectional dimension in finite samples does not constitute a constraint on feasibility even if the number of panels is much larger than the time dimension. Since the estimator is based on an asymptotic theory, one should be somewhat cautious when using large N and small T.

Moreover, very small T may be unsuitable for GETS because of the use of lagged changes of dependent and independent variables. The serial correlation (heteroscedasticity) may be addressed by employing the Wooldridge (2002) (modified Wald test for groupwise heteroscedasticity, see Greene, 2000) test. In the case of RE model, the Breusch–Pagan LM test may be used to explore the suitability of the RE model. While serial correlation is dealt with the Baltagi and Li (1991, 1995) test, robust standard errors are generated to solve the heteroscedasticity. However, CSD remains a caveat in the RE model because there is no test available for the case where $T > N$.

4. Empirical results

4.1. Panel unit root test results without structural breaks

We first apply Hadri (2000) panel unit root test. Table 2 presents the results for *ITY* and *STY*. The results indicate that the stationarity null can be rejected at the 1% statistical level, except for *ITY* (*STY*) in EU (CARTAGENA) samples. For the first differences of the two series, results reveal that they are stationary. However, these results can be misleading if there exist any misspecification error in the deterministic components (see Carrion-i-Silvestre et al., 2005). This misspecification error could be due to not considering for the presence of structural breaks. Further, this test does not allow for the presence of CSD. It is therefore vital to apply a panel unit root test that considers structural breaks and CSD.

4.2. Panel unit root test results with structural breaks

Carrion-i-Silvestre et al. (2005) modified the Hadri (2000) test to allow for multiple breaks by including dummy variables in the deterministic component of the model. This test allow for structural breaks in both the level and the slope of the variables. Given the number of time observations is not drastically large i.e. $T = 49$, it is reasonable to allow for a maximum of 2 structural breaks ($m^{max} = 2$). The modified Schwarz information criterion is utilized to estimate number of structural breaks associated with each individual.

Due to large number of countries under investigation, we did not report the country-by-country test results, however, we report only the country-specific break dates in Appendix 1 (Table A.1). We found that stationarity null hypothesis cannot be

Table 3
Carrion-i-Silvestre et al. (2005) stationarity test 1960–2012.

	Test statistic	*ITY* Bootstrap distribution (%)			Test statistic	*STY* Bootstrap distribution (%)		
		90	95	99		90	95	99
AFTA								
LM(λ)-*HOM*	11.220	5.293	8.119	10.028	10.295	14.941	17.005	20.990
LM(λ)-*HET*	3.271	7.270	7.830	8.274	8.370	11.830	16.022	19.227
EU								
LM(λ)-*HOM*	4.821	6.025	8.540	9.256	9.400	15.635	17.029	21.640
LM(λ)-*HET*	2.549	4.551	7.380	10.046	3.271	5.400	7.730	8.050
EFTA								
LM(λ)-*HOM*	13.480	15.272	17.880	17.953	3.941	4.398	5.634	6.750
LM(λ)-*HET*	10.004	13.265	15.028	18.920	6.300	8.341	10.925	13.260
CARTAGENA								
LM(λ)-*HOM*	1.290	3.442	5.403	6.038	3.291	18.901	19.000	22.375
LM(λ)-*HET*	3.477	5.460	6.940	7.995	6.938	19.045	20.011	21.048
NAFTA								
LM(λ)-*HOM*	2.311	3.219	3.500	4.038	4.190	6.273	10.290	12.443
LM(λ)-*HET*	1.036	2.377	5.058	6.285	2.266	4.290	7.127	11.294
MERCOSUR								
LM(λ)-*HOM*	8.311	11.202	12.114	15.922	7.295	8.003	13.640	18.091
LM(λ)-*HET*	7.125	9.035	11.277	11.924	2.113	5.162	8.299	11.976

Note: In the computation of the finite sample critical values, 2500 replications are used.

rejected for majority of the cases at the 5% level.[4] Some exceptions are Malta, Sweden and Philippines for *ITY* and Hungary, Peru, Uruguay and Poland for *STY*. An important aspect of this test is that in most countries, *ITY* and *STY* are affected by multiple breaks, i.e. two breaks (see Table A.1). Most break dates are located in the 1980s and 1990s and are plausible because they correspond to some major economic incidents that took place in these countries, for example economic reforms, integrations, oil shocks, etc. In some cases, the break dates are located from 2006 to 2008 and they highlight the presence of the recent global financial crisis.

Individual information can be pooled to derive the panel data stationary test results; Table 3 displays the results in which the long run variance are treated as both homogeneous and heterogeneous. The dependence of the test on the dates of the break is denoted by λ. We address the CSD in the data by generating the bootstrap distribution of the tests. Results revealed that the stationarity hypothesis cannot be rejected for both variables *ITY* and *STY* at the 5% level, except for *ITY* in AFTA sample.[5] Thus, based on these findings we infer that *ITY* and *STY* are characterized as stationary processes evolving around a broken trend.

4.3. GETS parsimonious estimates

The GETS procedure is utilized irrespective whether the series are $I(1)$ or $I(0)$. The procedure entails estimating a general unrestricted model (ARDL) using the FE and RE estimators. In doing so, *ITY* is regressed on the error correction variables viz., $[ITY_{it-1} - (\alpha + \beta STY_{it-1})]$ and on the lagged changes of *ITY* and *STY*. For NAFTA and MERCOSUR samples, we start with 4 lags because lags beyond 4 periods were statistically insignificant at the conventional levels. In the case of AFTA, EU, EFTA and CARTAGENA samples, lags up to 3 periods were statistically significant. Since there is no standard procedure to select the lag lengths in GETS procedure, it is therefore pragmatic to commence with a large lag structure and search for optimal lag order.

In the next stage, the general unrestricted models are reduced to manageable parsimonious versions by applying the standard variable deletion tests.[6] In these parsimonious versions, all estimates must be statistically significant at the conventional levels. Table 4 displays the parsimonious versions of FE and RE models. Robust standard errors are generated to eradicate the heteroscedasticity (in RE model estimates) and CSD (in FE model estimates) problems. Results show that all the coefficients are statistically significant at the conventional levels. The saving retention estimate indicates that 1% increase in *STY* leads to an increase in *ITY* by around 0.3% in EU and NAFTA countries. For AFTA and MERCOSUR countries, a 1% rise in *STY* leads to an increase in *ITY* by around 0.4% to 0.5%, respectively. The estimates for ETFA and CARTAGENA samples indicate that a 1% increase in *STY* will result in about 0.2% increase in *ITY* on average.

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[4] In the computation of the finite sample critical values, 2500 replications are used.
[5] Exceptions relates to the homogenous test.
[6] *PcGets* and *PcGive* software may be used to achieve the parsimonious GETS model.

Table 4
GETS parsimonious estimates with FE and RE models.

$$\Delta ITY_{it} = \lambda[ITY_{it-1} - (\alpha + \beta STY_{it-1})] + \sum\nolimits_{i=0}^{m} \gamma_i \Delta STY_{it-i} + \sum\nolimits_{i=1}^{n} \Phi_i \Delta ITY_{it-i} + u_{it}$$

	AFTA		EU		EFTA		CARTAGENA		MERCOSUR		NAFTA	
	FE	RE	FE	RE	FE	RE	FE	RE	FE	RE	FE	RE
Intercept	20.558	2.048	1.288	9.200	13.152	18.024	2.016	5.002	6.003	3.024	12.271	4.380
	(3.70)***	(0.37)***	(0.12)***	(1.11)***	(2.02)***	(1.50)***	(0.15)***	(0.52)***	(1.41)***	(0.25)***	(7.67)*	(0.63)***
λ	−0.401	−0.392	−0.211	−0.207	−0.534	−0.528	−0.406	−0.420	−0.196	−0.226	−0.381	−0.383
	(0.02)***	(0.04)***	(0.06)***	(0.05)***	(0.01)***	(0.05)**	(0.06)***	(0.08)**	(0.05)***	(0.07)***	(0.04)**	(0.02)***
βSTY_{it-1}	0.446	0.485	0.340	0.295	0.220	0.209	0.210	0.248	0.471	0.450	0.334	0.316
	(0.10)***	(0.14)***	(0.06)***	(0.12)**	(0.03)***	(0.06)***	(0.06)***	(0.12)**	(0.08)***	(0.19)**	(0.05)***	(0.04)***
$\gamma_1 \Delta STY_{it}$	0.366	0.359	–	0.170	0.003	–	0.625	–	0.114	0.281	–	0.009
	(0.03)***	(0.03)***		(0.06)**	(0.001)**		(0.18)***		(0.03)***	(0.09)***		(0.004)**
$\gamma_2 \Delta STY_{it-1}$	–	0.094	–	0.921	–	–	–	–	–	1.225	–	–
		(0.03)***		(0.10)***						(0.63)*		
$\gamma_3 \Delta STY_{it-2}$	0.084	–	1.281	–	–	0.920	1.381	0.720	–	–	0.911	–
	(0.03)**		(0.39)***			(0.22)***	(0.78)*	(0.30)**			(0.43)**	
$\gamma_4 \Delta STY_{it-3}$	0.128	0.060	–	–	2.009	–	0.100	–	0.390	–	2.070	–
	(0.03)***	(0.03)**			(1.19)*		(0.04)**		(0.04)***		(1.17)*	
$\gamma_5 \Delta STY_{it-4}$	–	–	–	–	–	–	–	–	0.625	–	1.290	–
									(0.30)**		(0.32)***	
$\Phi_1 \Delta ITY_{it-1}$	0.088	–	–	−0.005	–	–	–	–	–	−1.804	–	–
	(0.03)***			(0.001)**						(0.56)***		
$\Phi_2 \Delta ITY_{it-2}$	–	−0.112	0.820	–	–	–	–	0.127	−2.026	–	–	0.670
		(0.03)***	(0.46)*					(0.07)*	(1.13)*			(0.33)**
$\Phi_3 \Delta ITY_{it-3}$	0.082	–	−0.625	0.220	0.270	−2.301	0.530	1.320	0.162	–	–	–
	(0.02)***		(0.21)**	(0.04)***	(0.12)**	(1.28)*	(0.25)**	(0.49)**	(0.05)***			
$\Phi_4 \Delta ITY_{it-4}$	–	–	–	–	–	–	–	–	1.028	0.827	–	–
									(0.20)***	(0.24)***		

Notes: λ is the one period lagged error correction term. Robust standard errors are reported in parentheses to avoid heteroscedasticity (RE model). In the FE models, we robust the standard errors by applying Driscoll and Kraay (1998) procedure. The sample size is 1960–2012 in all cases. All regressions include a dummy variable; estimates not reported for brevity. The dummy variable is identified in the last column of Table A.1 in Appendix. Stata 12.0 is used to estimate the equations.

*** Significance at 1% level.

** Significance at 5% level.

* Significance 10% level.

The estimates of one period lagged error correction term (speed of adjustment) are around −0.4 (AFTA, CARTAGENA and NAFTA countries), −0.2 (EU and MERCOSUR countries) and −0.5 (EFTA countries). This implies the presence of negative feedback mechanism and in particular, if there are departures from equilibrium in the previous period, this departure is reduced by about 40% (AFTA, CARTAGENA and NAFTA countries), 20% (EU and MERCOSUR countries) and 50% (EFTA countries) in the current period. The speed of adjustment is lowest (highest) in the EU and MERCOSUR (EFTA) countries and this implies that the transition towards the equilibrium value will be sluggish (rapid). The disparity between the actual and equilibrium values of *ITY* and *STY* is attributed to shocks in these countries. The short-run effects of *STY* on *ITY* are generally positive. These impacts are transitory and may be unsustainable after some time.

To determine whether the FE or RE model is preferred, we applied the Hausman test. The null hypothesis is that the preferred model is random effects against the alternative of fixed effects. Estimates in Table 4 are subjected to Hausman test and the results reveal that the null hypothesis (preferred model is RE) can be rejected at the 1% level in all cases. The Hausman test results are as follows: $\chi^2(7) = 46.13$ with the p-value 0.00 (AFTA countries), $\chi^2(5) = 380.96$ with the p-value 0.00 (EU countries), $\chi^2(7) = 212.02$ with the p-value 0.00 (EFTA countries), $\chi^2(6) = 345.75$ with the p-value 0.00 (CARTAGENA countries), $\chi^2(10) = 147.64$ with the p-value 0.00 (MERCOSUR countries) and $\chi^2(5) = 144.02$ with the p-value 0.00 (NAFTA countries). In the light of these results, we shall argue that our preferred model is the FE and hence we shall follow this model in the rest of this paper.

Moreover, we shall investigate the impacts of six crucial agreements (AFTA, EU, EFTA, CARTAGENA, MERCOSUR and NAFTA) which may have contributed to increased capital mobility. To show the impacts of these international agreements, we construct sub-samples as follows: AFTA (pre-integration 1960–1991 and post-integration 1992–2012), EU (pre-integration 1960–1992 and post-integration 1993–2009), EFTA (post-integration 1960–2012),[7] CARTAGENA (post-integration 1970–2012), MERCOSUR (pre-integration 1960–1990 and post-integration 1991–2012) and NAFTA (pre-integration 1960–1993 and post-integration 1994–2012). We estimate the saving retention coefficient for above sub-samples. Our expectation is that the estimates of saving retention will decline in the second set of sub-samples to show an increase in the capital mobility. Table 5 presents the sub-sample estimates of β. In all cases, the results show that the estimate of saving retention has declined in the post-integration period, thus implying that international mobility of capital has increased in these countries.[8] Furthermore, the speed of adjustment in these countries has increased over-time, implying that in the presence of shocks, equilibrium errors are corrected quite rapidly.

The GETS procedure does not account for structural breaks and this may be a potential limitation if there exist statistically significant structural breaks in the data. Our unit root test results revealed that the *ITY* and *STY* series are characterized by two breaks, respectively (see Table A.1 in Appendix). To allow for breaks in the GETS procedure, we integrate a dummy variable in the FE and RE regressions. The dummy variable accounts for the level shifts that may have been caused due to the enforcement of integration agreements in these countries. Given that most break dates are consistent with the start date of the investment agreement, we therefore construct dummies to capture the impact of the event.

For AFTA sample, we used two dummy variables i.e. one to capture the impact of emergence of investment agreement and another to capture the effects of the Asian financial crisis that was experienced in the 1997–98 period. Moreover, the break dates in some countries are observed in the 2007 and 2008 period, thus highlighting the experience of the global financial crisis. In order to address this structural break, we estimated all samples excluding the global financial crisis period (i.e. 1960–2006) and achieved results that are qualitatively consistent to the results in Table 4.

4.4. Endogeneity tests

We utilize a simple procedure to identify the existence of endogeneity in *STY*. The growth in *STY* is regressed on the one period lagged ECM normalized on *ITY* (λECM_{it-1}). If *STY* is endogenous then the one period lagged ECM should be statistically significant with correct negative sign. For the sake of conformity in the results, we also regress the growth in *STY* on the one period lagged ECM normalized on *STY* ($\lambda ECMsty_{it-1}$). Table 6 presents these results. All the samples are estimated with the FE model. Results show that neither λECM_{it-1} nor $\lambda ECMsty_{it-1}$ are statistically significant with correct negative sign at the 5% level. In the AFTA countries sample, λECM_{it-1} is statistically significant at only 10% level. Overall, we infer that *STY* is not endogenous at the 5% statistical significance level and therefore there exist no endogeneity bias in our estimates. Moreover, introducing short-run dynamic variables in the equations did not change the overall results.

4.5. Cross-sectional dependence and other diagnostic tests

Our preferred estimates in Table 4 are tested for CSD, heteroscedasticity and serial correlation. The existence of CSD is not uncommon in time series panel data and if it is not dealt with, one may get little improvement in efficiency from panel estimators relative to a single time-series. There are many sources of CSD, for instance, spatial spillovers, interaction effects through trade or integration agreements and other common unobserved factors that influence all groups. Heteroscedasticity

[7] For EFTA and CARTAGENA, we do not have sufficient observations to conduct pre-integration estimations.

[8] The saving retention estimate for EFTA and CARTAGENA is low as 0.2 and we assume the estimate in the pre-integration period may be a bit high in magnitude.

Table 5
Integration impact on international capital mobility – GETS FE estimates.

$$\Delta ITY_{it} = \lambda[ITY_{it-1} - (\alpha + \beta STY_{it-1})] + \sum_{i=0}^{m} \gamma_i \Delta STY_{it-i} + \sum_{i=1}^{n} \Phi_i \Delta ITY_{it-i} + u_{it}$$

	AFTA		EU		CARTAGENA	EFTA	MERCOSUR		NAFTA	
	Pre-integ. 1960–1991	Post-integ. 1992–2012	Pre-integ. 1960–1992	Post-integ. 1993–2012	Post-integ. 1970–2012	Post-integ. 1960–2012	Pre-integ. 1960–1990	Post-integ. 1991–2012	Pre-integ. 1960–1993	Post-integ. 1994–2012
Intercept	2.387 (1.12)**	1.201 (0.37)***	14.294 (0.90)***	6.192 (2.50)**	10.382 (2.93)***	13.152 (2.02)***	1.293 (0.08)***	2.109 (0.48)***	8.281 (2.10)***	11.200 (2.85)***
λ	−0.339 (0.09)***	−0.510 (0.18)**	−0.368 (0.10)***	−0.551 (0.14)***	−0.449 (0.03)***	−0.534 (0.01)***	−0.380 (0.17)**	−0.404 (0.08)***	−0.350 (0.17)**	−0.372 (0.08)***
βSTY_{it-1}	0.540 (0.14)***	0.205 (0.06)***	0.426 (0.13)***	0.118 (0.03)***	0.205 (0.08)**	0.220 (0.03)***	0.601 (0.18)***	0.240 (0.08)**	0.501 (0.14)***	0.198 (0.05)***
$\gamma_1 \Delta STY_{it}$	–	0.095 (0.04)**	1.260 (0.38)***	0.821 (0.31)**	–	0.003 (0.001)**	–	0.785 (0.20)***	0.170 (0.09)*	–
$\gamma_2 \Delta STY_{it-1}$	0.092 (0.02)***	–	–	–	–	–	0.243 (0.05)***	0.674 (0.27)**	–	0.122 (0.04)**
$\gamma_3 \Delta STY_{it-2}$	–	–	–	–	0.437 (0.15)**	–	–	–	–	–
$\gamma_4 \Delta STY_{it-3}$	–	–	–	–	–	2.009 (1.19)*	–	0.920 (0.34)**	–	0.993 (0.30)***
$\Phi_1 \Delta ITY_{it-1}$	–	–	−0.877 (0.30)**	–	–	–	−0.837 (0.30)**	–	−1.284 (0.52)**	–
$\Phi_2 \Delta ITY_{it-2}$	−1.547 (0.53)**	−1.013 (0.46)**	–	−0.026 (0.01)**	−0.674 (0.20)***	–	–	–	–	–
$\Phi_3 \Delta ITY_{it-3}$	–	–	–	–	−2.736 (1.30)**	0.270 (0.12)**	–	–	–	–

Notes: Pre-integ. = pre-integration period. Post-integ. = post-integration period. λ is the one period lagged error correction term. We robust the standard errors (reported in parentheses) by applying Driscoll and Kraay (1998) procedure. Stata 12.0 is used to estimate the equations.

*** Significance at 1% level.

** Significance at 5% level.

* Significance 10% level.

Table 6
Endogeneity test for domestic saving 1960–2012.

	AFTA	EU	EFTA	CARTAGENA	MERCOSUR	NAFTA
$\Delta STY_{it} = \alpha + \lambda ECM_{it-1}$						
Intercept	15.720	25.003	4.393	14.500	−4.761	18.106
	(2.11)***	(2.17)***	(0.88)***	(3.44)***	(0.50)***	(3.08)***
λECM_{it-1}	−0.361	0.922	0.170	−0.635	0.140	1.205
	(0.22)*	(0.74)	(0.11)	(0.65)	(0.09)	(0.83)
$\Delta STY_{it} = \alpha + \lambda ECMsty_{it-1}$						
Intercept	2.645	12.037	−6.050	26.200	7.403	16.239
	(0.59)***	(1.41)***	(0.42)***	(2.48)***	(2.63)**	(2.48)***
$\lambda ECMsty_{it-1}$	0.824	−1.223	−0.230	−0.197	0.460	0.082
	(2.06)	(0.81)	(0.19)	(0.25)	(0.31)	(0.06)

Notes: Standard errors are reported in parentheses. $ECMsty_{it-1}$ and ECM_{it-1} are the one period lagged error correction term normalized on *STY* and *ITY*, respectively. Stata 12.0 is used to estimate these equations.
*** Significance at 1% level.
** Significance at 5% level.
* Significance 10% level.

Table 7
Diagnostic test results.

		Breusch–Pagan LM test of independence: χ^2 (*p*-value)	Modified Wald test for groupwise heteroscedasticity: χ^2 (*p*-value)	Wooldridge test for serial correlation: *F* (*p*-value)
AFTA	FE estimates	1233.85	13.470	0.211
		(0.00)***	(0.46)	(0.66)
EU	FE estimates	880.88	29.021	0.270
		(0.00)***	(0.30)	(0.65)
EFTA	FE estimates	583.04	47.206	10.528
		(0.04)**	(0.119)	(0.20)
CARTAGENA	FE estimates	609.25	18.279	3.005
		(0.00)***	(0.41)	(0.53)
MERCOSUR	FE estimates	918.11	104.240	32.100
		(0.01)***	(0.106)	(0.12)
NAFTA	RE estimates	897.02	18.270	27.372
		(0.00)***	(0.40)	(0.14)

Notes: χ^2 means chi-square. Wooldridge test for serial correlation produces an *F*-statistic with *p*-value.
*** Significance at 1% level.
** Significance at 5% level.

and serial correlation are also major issues that must be addressed. Table 7 displays the diagnostic test results. To identify the existence of CSD in the errors of the FE models, the Breusch–Pagan LM test of independence is used. In all cases, the null that residuals across entities are independent is rejected at the 1% level, except for EFTA countries at the 5% level. Moreover, we test for the presence of heteroscedasticity and first order serial correlation in our FE models. The modified Wald test for groupwise heteroscedasticity (Wooldridge test for serial correlation) fails to reject the null hypothesis of homoscedasticity (no serial correlation). With these findings, we assert that our estimates do not suffer from heteroscedasticity and serial correlation.

As our results strongly indicate the presence of common factors affecting the cross-sectional units, we therefore robust the standard errors by applying Driscoll and Kraay (1998) procedure. The standard errors reported in Tables 4 and 5 are subject to Driscoll and Kraay (1998) procedure. With these findings, we infer that our preferred GETS estimates are robust.

4.6. Heterogeneity

Since the classical FE estimator is homogeneous, it is therefore vital to verify the estimates with some heterogeneous panel estimators. It is well known that imposing homogeneity can cause the estimates to be biased if the true parameters are heterogeneous across countries. In addition to the application of FE estimator, we employ the dynamic fixed effects (DFE), mean group (MG) and pooled mean group (PMG) heterogeneous estimators proposed by Pesaran et al. (1999) to estimate the *ITY* and *STY* relationship for the ‘entire sample’. The latter three estimators require large *N* and large *T* and hence we combine all 44 countries and refer this as the ‘entire sample’. Table 8 presents these results.

The DFE, MG and PMG heterogeneous estimators yield estimates that are consistent to the FE (homogenous) estimates. The one period lagged ECM is around −0.3 and statistically significant at the 1% level. The saving retention coefficient is between 0.36 and 0.45, implying that 1% increase in *STY* leads to an increase in *ITY* by 0.36–0.45% on average. The short-run dynamic estimates are generally significant at the conventional levels.

Table 8
Estimates of saving retention using alternative methods.

$$\Delta ITY_{it} = \lambda[ITY_{it-1} - (\alpha + \beta STY_{it-1})] + \sum\nolimits_{i=0}^{m} \gamma_i \Delta STY_{it-i} + \sum\nolimits_{i=1}^{n} \Phi_i \Delta ITY_{it-i} + u_{it}$$

	FE (homogenous)	DFE (heterogeneous)	MG (heterogeneous)	PMG (heterogeneous)
Intercept	6.380 (1.20)***	2.300 (0.40)***	0.946 (0.30)**	5.044 (0.51)***
λ	−0.325 (0.10)***	−0.319 (0.09)***	−0.287 (0.05)***	−0.340 (0.10)***
β ln STY_{it-1}	0.401 (0.19)**	0.449 (0.01)***	0.360 (0.09)***	0.431 (0.08)***
$\gamma_1 \Delta \ln STY_{it}$	0.002 (0.001)**	–	–	–
$\gamma_2 \Delta \ln STY_{it-1}$	–	–	−0.006 (0.003)**	–
$\gamma_4 \Delta \ln STY_{it-3}$	–	0.016 (0.009)*	–	–
$\Phi_2 \Delta \ln ITY_{it-2}$	–	–	–	−0.602 (0.20)**

Notes: λ is the one period lagged error correction term. Robust standard errors are reported in the parentheses. The sample size is 1960–2012. Stata 12.0 is used to estimate the equations.
*** Significance at 1% level.
** Significance at 5% level.
* Significance 10% level.

4.7. Integration, financial intermediation and productivity

Over the last few decades, globalization and economic integration have become increasingly the dominant features of the world economy. Our results show that the international mobility of capital have exploded due to emergence of economic integration. In the light of these results, the question of whether economic integration affects financial intermediation and productivity may be of interest to the policy makers. Despite the rapidly growing literature on integration and capital mobility, there have been fewer attempts to explore the relationships between integration, financial intermediaries and productivity. Much of the empirical literature have focused on financial development and economic growth, for instance see Hasan et al. (2009), Kendall (2012), Bordo and Rousseau (2012), Deidda and Fattouh (2008), Levine (1999) and Tadesse (2002).

Following King and Levine (1993), we construct four measures of financial intermediation and study their behaviour in the pre- and post-integration periods and also investigate extent to which they are linked to real productivity. The first two indicators are the money demand indicators (*M1Y* and *M2Y*). *M1Y* is the ratio of a country's currency and demand deposits to its GDP. This indicator captures the scale of domestic currency funds held by individuals and corporations primarily for transactions. *M2Y* is the ratio of a country's broad money (sum of currency outside banks; demand deposits other than those of the central government; the time, savings, and foreign currency deposits of resident sectors other than the central government; bank and traveler's checks; and other securities such as certificates of deposit and commercial paper) held by agents to its GDP. The third indicator is the financial indicator (*LLY*), which measures the monetary and non-monetary liquid assets held by individuals. The final indicator is the ratio of a country's private credit to its GDP (*DCPSY*), where private credit is the credit extended to the private sector by commercial banks. This ratio implies the importance of the part played by the financial sector, especially the deposit money banks, in the financing of the economy.

In order to investigate the impacts of integration on financial intermediation and productivity, we adopt three ways to conduct this analysis. First, we compute the panel averages of *M1Y*, *M2Y*, *LLY* and *DCPSY*. Second, we estimate panel correlation ratios between the financial intermediation indicators and real productivity (y is measured as real GDP divided by total number of workers). Lastly, using Dumitrescu and Hurlin's (2012) panel Granger non-causality test, we perform causality analyses between the financial intermediation indicators and real productivity. Details about Dumitrescu and Hurlin (2012) test are available in Appendix 1. In all cases, we attain results for pre- and post-integration periods.[9] Table 9 presents these results. Results reveal that the average of *M1Y* (*M2Y*) has declined (increased) in the post-integration period in all panels. This is not an unexpected finding. Rao and Kumar (2011) find that the demand for *M1* in the U.S. declined over-time due to financial reforms. It is likely that economic integrations stimulate the economies of scale of money demand i.e. improves the financial system. The rise in *M2Y* may be due to the substitution effect (see Kumar, 2014). As income grows, individuals will economize more on cash (i.e. narrow money) and substitute with the check/savings accounts (i.e. broader aggregates).

The average correlation between the financial intermediation indicators and real productivity has increased in almost all panels in the post-integration period. The only exception is *M1Y*, where the correlation estimates have declined, albeit

[9] Except for EFTA and CARTAGENA where pre-integration results are not available due to data limitations.

Table 9
Financial intermediation in pre- and post-integration.

	AFTA		EU		EFTA CARTAGENA		MERCOSUR		NAFTA	
	Pre-integ. 1960–1991	Post-integ. 1992–2012	Pre-integ. 1960–1992	Post-integ. 1993–2012	Post-integ. 1960–2012	Post-integ. 1970–2012	Pre-integ. 1960–1990	Post-integ. 1991–2012	Pre-integ. 1960–1993	Post-integ. 1994–2012
M1Y	0.395	0.140	0.287	0.191	0.172	0.255	0.309	0.184	0.302	0.214
M2Y	0.250	0.399	0.361	0.445	0.302	0.348	0.197	0.320	0.274	0.395
LLY	0.300	0.472	0.279	0.385	0.373	0.481	0.256	0.411	0.309	0.490
DCPSY	0.287	0.503	0.315	0.482	0.290	0.256	0.384	0.470	0.351	0.427
Corr(*M1Y*, *y*)	0.839	0.772	0.925	0.920	0.711	0.699	0.850	0.799	0.786	0.695
Corr(*LLY*, *y*)	0.892	0.999	0.784	0.946	0.999	0.945	0.761	0.892	0.910	0.999
Corr(*DCPSY*, *y*)	0.533	0.840	0.872	0.993	0.830	0.991	0.745	0.890	0.742	0.995
Corr(*M2Y*, *y*)	0.704	0.996	0.825	0.990	0.974	0.860	0.811	0.985	0.826	0.999
Panel causality:	$M1Y \rightarrow y$	$M1Y \rightarrow y$	$M1Y \rightarrow y$	$M1Y \rightarrow y$	$M1Y \rightarrow y$	$M1Y \rightarrow y$	$M1Y \rightarrow y$	$M1Y \rightarrow y$	$M1Y \rightarrow y$	$M1Y \rightarrow y$
M1Y and *y*	6.38 (0.06)***	17.29 (0.00)*	14.19 (0.00)*	38.27 (0.00)*	23.60 (0.00)*	22.36 (0.00)*	34.84 (0.00)*	34.09 (0.00)*	5.37 (0.08)***	26.34 (0.00)*
	$M1Y \leftarrow y$	$M1Y \leftarrow y$	$M1Y \leftarrow y$	$M1Y \leftarrow y$	$M1Y \leftarrow y$	$M1Y \leftarrow y$	$M1Y \leftarrow y$	$M1Y \leftarrow y$	$M1Y \leftarrow y$	$M1Y \leftarrow y$
	11.20 (0.00)*	8.26 (0.00)*	11.27 (0.00)*	12.09 (0.00)*	14.35 (0.00)*	6.32 (0.03)**	22.48 (0.00)*	14.86 (0.00)*	15.06 (0.00)*	36.01 (0.00)*
Panel causality:	$LLY \rightarrow y$	$LLY \rightarrow y$	$LLY \rightarrow y$	$LLY \rightarrow y$	$LLY \rightarrow y$	$LLY \rightarrow y$	$LLY \rightarrow y$	$LLY \rightarrow y$	$LLY \rightarrow y$	$LLY \rightarrow y$
LLY and *y*	2.82 (0.69)	27.05 (0.00)*	5.01 (0.09)***	45.19 (0.00)*	14.73 (0.00)*	31.07 (0.00)*	4.65 (0.28)	17.04 (0.00)*	2.51 (0.70)	41.94 (0.00)*
	$LLY \leftarrow y$	$LLY \leftarrow y$	$LLY \leftarrow y$	$LLY \leftarrow y$	$LLY \leftarrow y$	$LLY \leftarrow y$	$LLY \leftarrow y$	$LLY \leftarrow y$	$LLY \leftarrow y$	$LLY \leftarrow y$
	1.36 (0.80)	15.42 (0.00)*	3.45 (0.41)	19.34 (0.00)*	21.89 (0.00)*	23.16 (0.00)*	5.14 (0.08)***	22.88 (0.00)*	4.16 (0.26)	25.19 (0.00)*
Panel causality:	$DCPSY \rightarrow y$	$DCPSY \rightarrow y$	$DCPSY \rightarrow y$	$DCPSY \rightarrow y$	$DCPSY \rightarrow y$	$DCPSY \rightarrow y$	$DCPSY \rightarrow y$	$DCPSY \rightarrow y$	$DCPSY \rightarrow y$	$DCPSY \rightarrow y$
DCPSY and *y*	3.80 (0.51)	42.38 (0.00)*	1.29 (0.85)	28.22 (0.00)*	14.17 (0.00)*	18.16 (0.00)*	3.72 (0.53)	13.95 (0.00)*	3.29 (0.48)	17.92 (0.00)*
	$DCPSY \leftarrow y$	$DCPSY \leftarrow y$	$DCPSY \leftarrow y$	$DCPSY \leftarrow y$	$DCPSY \leftarrow y$	$DCPSY \leftarrow y$	$DCPSY \leftarrow y$	$DCPSY \leftarrow y$	$DCPSY \leftarrow y$	$DCPSY \leftarrow y$
	5.22 (0.08)***	31.97 (0.00)*	2.01 (0.75)	10.24 (0.00)*	12.27 (0.00)*	34.01 (0.00)*	4.46 (0.24)	20.46 (0.00)*	5.53 (0.07)***	22.63 (0.00)*
Panel causality:	$M2Y \rightarrow y$	$M2Y \rightarrow y$	$M2Y \rightarrow y$	$M2Y \rightarrow y$	$M2Y \rightarrow y$	$M2Y \rightarrow y$	$M2Y \rightarrow y$	$M2Y \rightarrow y$	$M2Y \rightarrow y$	$M2Y \rightarrow y$
M2Y and *y*	1.29 (0.83)	15.78 (0.00)*	4.87 (0.18)	38.14 (0.00)*	18.72 (0.00)*	26.09 (0.00)*	3.54 (0.42)	18.24 (0.00)*	2.50 (0.70)	45.80 (0.00)*
	$M2Y \leftarrow y$	$M2Y \leftarrow y$	$M2Y \leftarrow y$	$M2Y \leftarrow y$	$M2Y \leftarrow y$	$M2Y \leftarrow y$	$M2Y \leftarrow y$	$M2Y \leftarrow y$	$M2Y \leftarrow y$	$M2Y \leftarrow y$
	4.80 (0.22)	11.45 (0.00)*	2.42 (0.71)	26.08 (0.00)*	21.94 (0.00)*	13.42 (0.00)*	4.92 (0.20)	15.73 (0.00)*	3.61 (0.48)	28.11 (0.00)*

Key: *M1Y* = narrow money (% of GDP); *LLY* = liquid liabilities (% of GDP); *DCPSY* = domestic credit to private sector by banks (% of GDP); *M2Y* = broad money (% of GDP); *y* = real productivity (real GDP/total number of workers).

Notes: Pre-integ. = pre-integration period. Post-integ. = post-integration period. For the panel causality tests, *F* test statistics and p-values are reported below the variables. *p*-Values are in parentheses. $X \rightarrow Y$ indicates the causality is running from *X* to *Y* and $X \leftarrow Y$ indicates the vice versa. Corr means correlation, for example, Corr(*M1Y*, *y*) indicates panel correlation between *M1Y* and *y*.

*** Significance at 1% level.

** Significance at 5% level.

* Significance 10% level.

slightly. Dumitrescu and Hurlin (2012) tests reveal the existence of bi-variate causality between *M1Y* and *y* in pre- and post-integration periods. For other variables (*M2Y*, *LLY* and *DCPSY*), there is lack of causality between the indicators and real productivity in the pre-integration period. However, the post-integration periods are characterized by bi-variate causality between the indicators and real productivity. There is evidence in the growth literature that *M2Y* contributes positively to output growth (Rao and Hassan, 2011). The positive relationship between *LLY* (and *DCPSY*) and *y* is supported by King and Levine (1993).

5. Conclusion

In this paper, we utilized the Feldstein–Horioka puzzle to investigate the impact of regional integration agreements (AFTA, EU, EFTA, CARTAGENA, MERCOSUR and NAFTA) on the international capital mobility. In doing so, we extend the univariate time series general to specific (GETS) method of Hendry (1995) to estimate the cointegrating equation and dynamic adjustments in panel data. Using the classical fixed and random effects estimators, we estimate the long- and short-run effects in the same model and we show that it is possible to estimate the lagged adjustment process which implies negative feedback mechanism. The procedure used is general enough to allow for the presence of endogeneity, heteroscedasticity, serial correlation and cross-sectional dependence in the residuals.

We find fairly similar degrees of international capital mobility in our samples. For EU and NAFTA countries, 1% increase in *STY* leads to an increase in *ITY* by around 0.3% on average. In the case of AFTA and MERCOSUR countries, a 1% rise in *STY* leads to an increase in *ITY* by around 0.4% to 0.5%, respectively. The estimates for ETFA and CARTAGENA samples indicate that a 1% increase in *STY* will result in about 0.2% increase in *ITY* on average. In all cases, the results show that the estimate of saving retention has declined in the post-integration period, thus implying that international mobility of capital has increased in these countries. Moreover, our ‘entire sample’ results are consistent to the results attained by applying Pesaran et al.’s (1999) dynamic fixed effects, mean group and pooled mean group heterogeneous estimators.

The disparity between the actual and equilibrium values of *ITY* and *STY* is attributed to shocks in these countries (among other reasons) and the estimated one period lagged error correction term attempts to explain the presence of negative feedback mechanism. In particular, if there are departures from equilibrium in the previous period, this departure is reduced by about 40% (AFTA, CARTAGENA and NAFTA countries), 20% (EU and MERCOSUR countries) and 50% (EFTA countries) in the current period. We find that in the post-integration period, the speed of adjustment has increased in these countries.

Moreover, our results shed light on the question whether economic integration affects financial intermediation and productivity. Following King and Levine (1993), we constructed four measures of financial intermediation (narrow money to GDP ratio (*M1Y*), broad money to GDP ratio (*M2Y*), monetary and non-monetary liquid assets held by individuals to GDP ratio (*LLY*) and private credit to GDP ratio (*DCPSY*)) and investigate their behaviour in the pre- and post-integration periods and also explore extent to which they are linked to real productivity. Results reveal that the average of *M1Y* (*M2Y*) has declined (increased) in the post-integration period in all panels. This implies that economic integrations stimulate the economies of scale of money demand. Furthermore, integrations yield a greater degree of substitution effect i.e. agents economize more on cash (narrow money) and substitute with the check/savings accounts (broader money). Application of Dumitrescu and Hurlin (2012) test revealed the existence of bi-variate causality between *M1Y* and *y* in pre- and post-integration periods. In the cases of *M2Y*, *LLY* and *DCPSY*, there is lack of causality between the indicators and real productivity in the pre-integration period. However, the post-integration periods yields bi-variate causality between the indicators and real productivity.

Needless to say, there are limitations in this paper. First, the fixed effects and random effects estimators assume homogeneity in the parameters and this can cause the estimates to be biased if the true parameters are heterogeneous across countries. It is therefore important to cross-check the estimates with some heterogeneous panel estimators. Second, we did not consider the presence of endogenous structural breaks and nonlinearity in the data. However, these issues are outside the scope of this current paper. We hope that these issues will gain attention in future work.

Appendix 1.

See Figs. A.1 and A.2 and Table A.1.

Dumitrescu and Hurlin (2012) panel Granger non-causality test

Dumitrescu and Hurlin (2012) propose Granger noncausality tests for heterogeneous panel data models. Consider *Y* is the dependent variable and *X* is the explanatory variable, then the model specification is:

$$\Delta \ln Y_{i,t} = \upsilon + \sum_{a=1}^{n} \theta_a \Delta \ln Y_{i,t-a} + \sum_{a=0}^{n} k_a \Delta \ln X_{i,t-a} + u_{i,t}$$

$$\Delta \ln X_{i,t} = \upsilon + \sum_{a=1}^{n} \theta_a \Delta \ln X_{i,t-a} + \sum_{a=0}^{n} k_a \Delta \ln Y_{i,t-a} + u_{it}$$

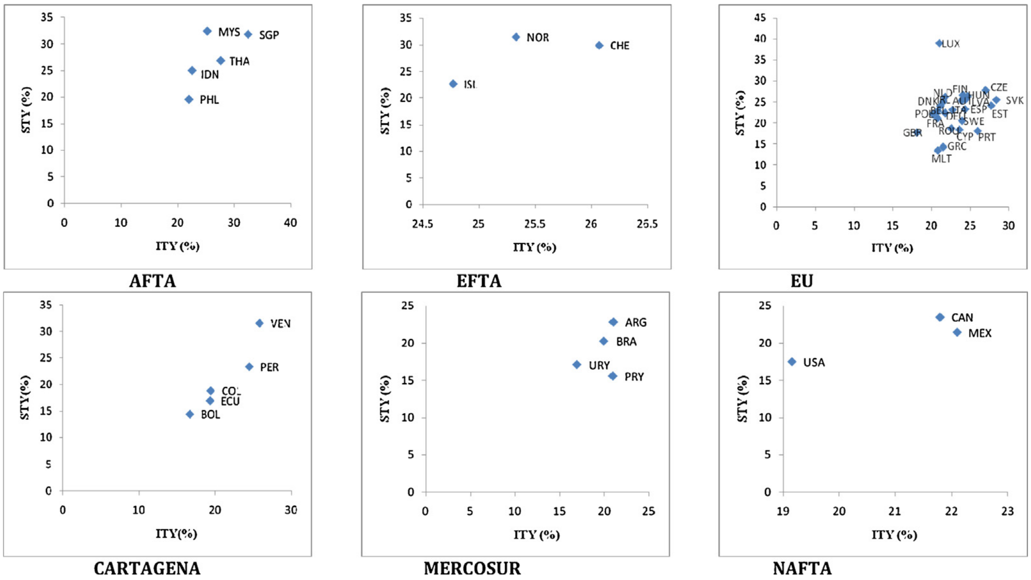

Fig. A.1. *ITY* and *STY* averages for each country.

where i is individual of the panel ($i = 1, \ldots, N$), t is time period ($t = 0, \ldots, T$) and n is the maximum number of considered lags. It is assumed that there are balanced panels and identical lag orders (a) for all cross section units. The F-tests are utilized to formulate inferences related to Granger non-causality. We test the following hypotheses:

$$H_0 : k_i^a = 0 \quad \forall i \in [1, N], \quad \forall a \in [0, n],$$
$$H_1 : k_i^a \neq 0 \quad \forall i \in [1, N], \quad \forall a \in [0, n],$$

If the null hypothesis is not rejected, this means that there exists no causality between the variables.

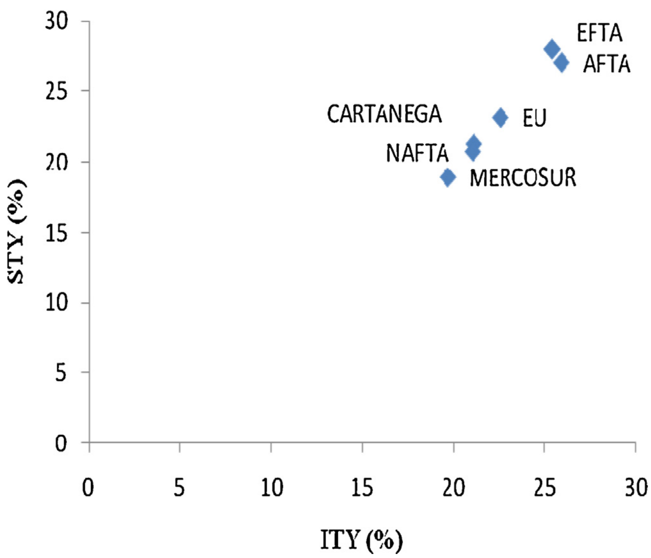

Fig. A.2. *ITY* and *STY* averages for each panel.

Table A.1
Carrion-i-Silvestre et al. (2005) test country-specific break dates.

Country	ITY		STY		Dummy variables created for GETS FE/RE model
	Break 1	Break 2	Break 1	Break 2	
AFTA					
Indonesia	1990	1992	1992	1997	DUM_{AFTA} DUM_{crisis}
Malaysia	1992	2006	1994	1998	DUM_{AFTA} DUM_{crisis}
Philippines	1992	1998	2006	1997	DUM_{AFTA} DUM_{crisis}
Singapore	1993	1997	1992	1999	DUM_{AFTA} DUM_{crisis}
Thailand	1991	1992	1993	1997	DUM_{AFTA} DUM_{crisis}
EU					
Austria	1992	1994	1993	2002	DUM_{EU}
Belgium	1993	1994	1995	2007	DUM_{EU}
Cyprus	1999	2006	1993	1995	DUM_{EU}
Czech Rep	2001	2002	1994	1997	DUM_{EU}
Denmark	1992	1995	2006	2007	DUM_{EU}
Estonia	1997	2005	1998	2002	DUM_{EU}
Finland	2007	2008	1987	1993	DUM_{EU}
France	1993	1994	1993	1994	DUM_{EU}
Germany	1993	1994	1995	2000	DUM_{EU}
Greece	1995	2000	1992	1994	DUM_{EU}
Hungary	2000	2006	2001	2005	DUM_{EU}
Ireland	1994	2007	1994	2007	DUM_{EU}
Italy	1993	2007	1994	1995	DUM_{EU}
Latvia	1994	1995	1993	1994	DUM_{EU}
Luxembourg	1992	1993	1996	1999	DUM_{EU}
Malta	2004	2005	1994	2004	DUM_{EU}
Netherlands	1988	1992	1993	2000	DUM_{EU}
Poland	1992	1993	1993	1994	DUM_{EU}
Portugal	1993	1996	1994	1995	DUM_{EU}
Romania	1994	1995	1992	1994	DUM_{EU}
Slovakia	1994	2005	1993	1995	DUM_{EU}
Spain	1995	2000	2001	2006	DUM_{EU}
Sweden	1993	1995	1995	2004	DUM_{EU}
UK	1994	2007	1995	1996	DUM_{EU}
EFTA					
Iceland	1985	1988	2001	2005	DUM_{EFTA}
Norway	2006	2007	2005	2007	DUM_{EFTA}
Switzerland	2007	2008	1990	1998	DUM_{EFTA}
CARTAGENA					
Bolivia	1987	1990	1986	1988	$DUM_{CARTAGENA}$
Columbia	1988	1989	1988	1990	$DUM_{CARTAGENA}$
Ecuador	2000	2005	1991	2007	$DUM_{CARTAGENA}$
Peru	1989	1990	1987	1989	$DUM_{CARTAGENA}$
Venezuela	1987	1988	1988	1989	$DUM_{CARTAGENA}$
MERCOSUR					
Argentina	1990	1991	1991	2000	$DUM_{MERCOSUR}$
Brazil	1991	1992	1993	1994	$DUM_{MERCOSUR}$
Paraguay	1991	2004	1991	2001	$DUM_{MERCOSUR}$
Uruguay	1992	1994	1991	1992	$DUM_{MERCOSUR}$
NAFTA					
Canada	1993	1995	1994	1996	DUM_{NAFTA}
Mexico	1994	1995	1994	2002	DUM_{NAFTA}
USA	1995	2007	1993	2007	DUM_{NAFTA}

Notes: DUM_{AFTA} is a dummy variable to capture the implementation of AFTA agreement. DUM_{AFTA} is 1 in 1992–94 and 0 otherwise. DUM_{crisis} captures that the Asian financial crisis in 1997–98. DUM_{crisis} is 1 in 1997–98 and 0 otherwise. DUM_{EU} is a dummy variable to capture the impacts of EU agreement. DUM_{AFTA} is 1 in 1993–95 and 0 otherwise. DUM_{EFTA} is a dummy variable to capture the impacts of EFTA agreement. DUM_{EFTA} is 1 in 1960–63 and 0 otherwise. $DUM_{CARTAGENA}$ is a dummy variable to capture the impacts of CARTAGENA agreement. $DUM_{CARTAGENA}$ is 1 in 1988–90 and 0 otherwise. $DUM_{MERCOSUR}$ is a dummy variable to capture the impacts of MERCOSUR agreement. $DUM_{MERCOSUR}$ is 1 in 1991–93 and 0 otherwise. DUM_{NAFTA} is a dummy variable to capture the impacts of NAFTA agreement. DUM_{NAFTA} is 1 in 1994–96 and 0 otherwise.

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