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RESEARCH 

# Capital mobility in Latin American and Caribbean countries: new evidence from dynamic common correlated effects panel data modeling

Vasudeva N. R. Murthy[1] and Natalya Ketenci[2*]

* Correspondence: nketenci@yeditepe.edu.tr
[2]Department of Economics, Yeditepe University, 26 Agustot Yerlesimi, Atasehir, 34755 Istanbul, Turkey
Full list of author information is available at the end of the article

## Abstract

This study investigates the degree of capital mobility in a panel of 16 Latin American and 4 Caribbean countries during 1960 to 2017 against the backdrop of the Feldstein-Horioka hypothesis by applying recent panel data techniques. This is the first study on capital mobility in Latin American and Caribbean countries to employ the recently developed panel data procedure of the dynamic common correlated effects modeling technique of Chudik and Pesaran (J Econ 188:393–420, 2015) and the error-correction testing of Gengenbach, Urbain, and Westerlund (Panel error correction testing with global stochastic trends, 2008, J Appl Econ 31:982–1004, 2016). These approaches address the serious panel data econometric issues of cross-section dependence, slope heterogeneity, nonstationarity, and endogeneity in a multifactor error-structure framework. The empirical findings of this study reveal a low average (mean) savings–retention coefficient for the panel as a whole and for most individual countries, as well as indicating a cointegration relationship between saving and investment ratios. The results indicate that there is a relatively high degree of capital mobility in the Latin American and Caribbean countries in the short run, while the long-run solvency condition is maintained, which is due to reduced frictions in goods and services markets causing increase competition. Increased capital mobility in these countries can promote economic growth and hasten the process of globalization by creating a conducive economic environment for FDI in these countries.

**Keywords:** Dynamic common correlated effects, Panel-error correction modeling, Cross-sectional dependence, Unobserved common factors, Slope heterogeneity, Capital mobility

**JEL classifications:** C10, C59, F20

## Introduction

Economists and policymakers have been studying the dynamic role of capital mobility in economic growth recently, especially in emerging countries in general and in Latin American and Caribbean countries in particular, which are experiencing large inflows of capital from abroad. One source is the recent unconventional monetary expansion in the United States through large-scale quantitative easing (QE), undertaken through extensive purchases of assets by its Federal Reserve. Lower U.S. interest rates and other phenomena, such as an appetite for increased global risk, improvements in these developing countries' macroeconomic fundamentals, and the rapid progress in information technology, have all contributed to the increase in capital flow to these economies.

Ford and Horioka (2016), in explaining the Feldstein-Horioka puzzle (Feldstein and Horioka 1980; Feldstein 1983), state that net transfers of capital among countries depend not only on the integration of financial markets but also on the integration of the goods and services markets. Ko and Funashima (2019) find evidence that large markets have higher correlations between savings and investments compared with mid- and small-sized countries. Eaton et al. (2016) present empirical evidence that financial friction in the goods and services markets reduces the degree of capital mobility. In fact, they contend that removing the frictions in the goods and services markets reduces considerably the dependence of domestic investment on domestic saving, leading to a greater degree of capital mobility in the observed Feldstein-Horioka puzzle, which is estimated by the following equation:

$$(I/Y)_{\mathrm{it}} = \alpha + \beta_{it}\left(S/Y\right)_{it} + \varepsilon_{it}$$

where *I/Y* is the investment ratio to gross domestic product (GDP) in country *i* and period *t*. S/Y is the saving ratio as a percentage of GDP and $\beta_{it}$ is the saving retention coefficient, which indicates the capital mobility level in country *i*. $\varepsilon_{it}$ indicates the error term of the regression model. In countries where capital mobility is high, the savings–retention coefficient is expected to be low reflecting a low level of correlation between domestic investments and savings. However, Feldstein and Horioka (1980) illustrate empirically that the correlation between investment and saving ratios is high in developed countries, where it is expected to be low, which has created the Feldstein-Horioka puzzle.

Eaton et al. (2016) argue, given that there is no "home-country bias" and an absence of financial friction in the goods markets and financial markets, in recipient countries, there are many benefits from an inflow of capital where there is a low level of domestic savings as well as crowding-out effects of a budget deficit. These factors have been instrumental in enabling a developing economy to pursue the most profitable investment opportunities and acquire foreign funds to finance domestic investment projects. This then limits the tax burden on relatively immobile domestic factors of production, smooths domestic consumption, and, ultimately, improves resource allocation and the economic welfare in the recipient countries (Obstfeld and Rogoff 2010; Boschi 2012; Ghosh et al. 2012; Ahmed and Zlate 2013; Koepke 2019; Zheng et al. 2019; Al-Jassar and Moosa 2020).

At the same time, an economy that is witnessing a steady capital inflow, despite some benefits, may experience many undesirable economic consequences and distortions, the most important being the inability to implement independent monetary policy actions. A review of capital inflow data reveals that many Latin American countries, unlike in the 1970s and 1980s, have experienced larger capital inflows recently. External shocks, like low-interest rates or economic slowdowns in developed countries, "push" investors to emerging markets, like Latin America, and are considered key factors in attracting foreign investments (Calvo et al. 1993; Fernandez-Arias 1996; Aizenman and Binici 2016; Kang and Kyunghun 2019; Koepke 2019; Eller et al. 2020). However, Baek (2006) argues that together with low international interest rates, which are push factors, the most important factor attracting foreign investment to Latin America is strong domestic economic growth, which is considered a pull factor for foreign investments. Fomina (2021), with the help of systematic and structural-logical techniques, modeled the stages forming pull factors that increase the competitive advantages of Latin America and improve its investment attractiveness.

To analyze country and regional international case studies in this context and derive current and future policy implications, a reliable and econometrically robust quantitative measure of the prevailing degree of capital mobility, such as an estimate of the savings–retention coefficient, is warranted. The main objectives of this study are to test whether capital mobility exists and to quantify its degree using the size of the savings–retention coefficient by investigating a panel of 16 Latin American and 4 Caribbean countries during the period 1960–2017.

To investigate capital mobility in Latina American and Caribbean countries, Murthy (2009) applies the panel group fully modified ordinary least squares (FM-OLS) estimator developed by Pedroni (2000, 2001) for the period 1960–2002. Our study employs panel data over a longer period (1960 to 2017) by using a recently developed robust panel data estimation technique, the dynamic common correlated effects mean group (DCCEMG) estimator from Chudik and Pesaran (2015). To test whether our results are robust, we also report the savings–retention coefficients estimated from applying another panel data estimator, Pesaran's (2006) common correlated effects mean group (CCEMG). Furthermore, to test whether there is a cointegrating relationship between the saving and investment ratios in the presence of cross-sectional dependence, we conduct the Gengenbach et al. (2008, 2016) error-correction tests at the panel and individual country levels. To the best of our knowledge, no other study in the literature has employed these panel data estimations along with a relatively long period sample to examine Latin American and Caribbean countries. The rest of the paper is organized as follows. Literature survey section briefly reviews the literature on capital mobility; Model specification and data section provides the specification of the employed model and the data; Empirical results section presents results of the empirical analysis conducted in this study and finally the last section concludes with the discussion of policy implications.

## Literature survey

A literature survey on capital mobility shows many studies investigating the prevalence of capital mobility using the Feldstein-Horioka hypothesis and framework in both developed and developing countries. As explained in the introduction, in the literature on capital mobility in international macroeconomics, Feldstein-Horioka hypothesis is a major puzzle. It should be noted that Feldstein and Horioka were the pioneers in statistically determining whether capital mobility exists in 16 Organization of Economic Cooperation and Development (OECD) countries during 1960 to 1974. As explained, using a simple regression model, they find some econometric evidence, contrary to the expected theoretical notion, that capital mobility is absent in this group of countries.

To test the presence of capital mobility in the 16 OECD countries, they specify an econometric model (FH-Model).

According to the FH-model, if the estimated savings–retention coefficient is not statistically different from zero in the estimated model, then there is perfect capital mobility. In contrast, if the value of the estimated coefficient $\beta$ is close to one and is statistically significant, then the evidence supports that capital is immobile. In the economic literature on capital mobility, the reasoning is that statistically significant lower values of $\beta$ denote the prevalence of a reasonable or moderate degree of capital mobility, although some development economists suggest a special cutoff savings–retention coefficient value of 0.60 (see, Murphy (1984)), especially for developing countries, implying the presence of a moderate degree of capital mobility. In their econometric study, Feldstein and Horioka (1980) find an estimated statistically significant savings–retention coefficient of 0.887, with a computed t-value of 12.67, and a coefficient of determination (R-square) for the model as a whole of 0.91. Therefore, they conclude that in these countries, contrary to the expected theoretical notion of capital mobility, the empirical evidence reflects the absence of capital mobility. Hence, the name for this unexpected phenomenon is the Feldstein-Horioka puzzle. Since the publication of their important research papers, many econometric studies have attempted to determine empirically the degree of capital mobility in various countries, in different groups of countries, over different time periods, using both time-series and panel data.

In addition, there are many studies dealing with the issue of capital mobility in OECD countries. Details of these studies are in Coakley et al. (1998), Apergis and Tsoumos (2009), and Singh (2016). Many economists, using cross-sectional, time-series, and panel data, applying various econometric techniques, have tested the celebrated Feldstein-Horioka puzzle (FH Puzzle) on the existence of varying degrees of capital mobility. Among these, notable studies dealing with the FH Puzzle using time series and cross-sectional data are Coakley and Kulasi (1997), Narayan (2005), Rocha (2006), Chen and Shen (2015), Ketenci (2012), Ma and Li (2016), Dash (2019), Ko and Funashima (2019), Zargar et al. (2019), Bineau (2020), and Akkoyunlu (2020). There are more studies that employ panel data; for example, Ho (2002), Holmes (2005), Kim et al. (2005), Murthy (2005, 2009), Payne and Kumazawa (2005), Murthy and Anoruo (2010), Narayan and Narayan (2010), Kumar and Rao (2011), Bangake and Eggoh (2012), Holmes and Otero (2014), Johnson and Lamdin (2014), Hernandez (2015), Bibi and Jalil (2016), Drakos et al. (2017), Pata (2018), and Eyuboglu and Uzar (2020). Our study combines estimations for both individual and panel series data.

Further, many studies have attempted to solve the FH puzzle for developed countries. However, as in our case, others have tested the FH puzzle in developing countries, where large investment inflows are essential factors for economic growth. As may be expected in studies in developing countries, Narayan (2005) and Rocha (2006) find a high saving–investment correlation, indicating restricted capital mobility in these countries. However, other studies find increasing capital mobility in developing countries; for example, Holmes (2005), Kim et al. (2005), Murthy (2005, 2009), Payne and Kumazawa (2005), Murthy and Anoruo (2010), Bangake and Eggoh (2012), Hernandez (2015), and Ketenci (2015). Further, there are several studies that have investigated quantitatively the impact of financial flows on the FH puzzle by modeling algorithms focusing on financial risk analysis; for example, Bai and Zhang (2010), Kou et al. (2014), Camarero et al. (2019), Causevic (2020), and Wang et al. (2020).

A review of these studies indicates that only a small number, with the exception of Murthy (2009), focus on the degree of capital mobility exclusively in Latin American countries. Holmes (2005), using data for 1979 to 2001, finds a savings–retention coefficient of 0.33, applying a FM-OLS estimator for a panel of 13 Latin American countries. Murthy (2009), conducting a panel cointegration analysis using data for 1960 to 2002, employing the Pedroni group FM-OLS estimator, reports a savings–retention coefficient of 0.46 with incorporated common time dummies and 0.48 without common time dummies for 14 Latin American countries and 5 Caribbean countries.

Most studies on capital mobility that apply panel data methods suffer from several econometric shortcomings. These studies, with the exceptions of Hernandez (2015) and Bibi and Jalil (2016), do not address some important econometric issues that plague panel data such as the existence of observed and unobserved common effects, cross-sectional dependence, parameter heterogeneity, and endogeneity in a multifactor dynamic error framework (Pesaran 2006; Eberhardt and Bond 2009; Cavalcanti et al. 2011; Pesaran and Tosetti 2011; Chudik and Pesaran 2015; Ditzen 2016). Hernandez (2015) and Bibi and Jalil (2016) employ Pesaran's (2006) static CCEMG. While Bibi and Jalil (2016) include a large panel consisting of 88 widely diverse countries. Hernandez (2015) uses panel data consisting of 18 emerging economies, looking at quarterly data from 2000Q1 to 2012Q4. Whereas Bibi and Jalil (2016) apply the static Pesaran CCEMG, Hernandez employs the augmented mean group (AMG) estimator (Eberhardt and Bond 2009), which controls for cross-sectional dependence in a static multifactor error structure framework. Eyuboglu and Uzar (2020) employ both the static CCEMG and the AMG estimators. The CCEMG estimator, although a robust panel data method, has shown that it does not yield consistent estimates in a dynamic multifactor error framework (Chudik and Pesaran 2015; Everaert and De Groote 2016; Ditzen 2016). Therefore, to overcome existing econometric shortcomings, we apply the DCCEMG estimation procedure to test whether capital mobility exists examining a panel of 20 Latin American and Caribbean countries 1960 to 2017. The choice of countries and the time period are dictated by the availability of reliable and complete data sets.

## Model specification and data

Our study tests the economic hypothesis of the FH puzzle. We do so by applying the DCCEMG estimator. The DCCEMG is a modified estimator for handling dynamic and heterogeneous coefficients of a panel model that incorporates lagged dependent and weakly exogenous regressors. Following Chudik and Pesaran (2015) and Ditzen (2016, 2018), the small sample time series bias is controlled by using the recursive correction method. Using similar notations of Ditzen (2016, 2018), we specify the model in a multifactor error structure framework, as shown in Eqs. (1)–(3) (Pesaran 2006; Chudik and Pesaran 2015; Baltagi 2015, 2020; Eberhardt and Teal 2011, 2013; Cavalcanti et al. 2011).

$$\left(\frac{I}{Y}\right)_{i,t} = \alpha_{1i} + d_{it} + \beta_i \left(\frac{S}{Y}\right)_{i,t} + \mu_{i,t} \tag{1}$$

$$\mu_{i,t} = \tau_i f_{\mathrm{t}} + \varepsilon_{i,t} \tag{2}$$

where $(I/Y)_{i,t}$ and $(S/Y)_{i,t}$ are the ratio of gross capital to GDP and the ratio of gross domestic savings to GDP, respectively. $\beta_i$ denotes the country-specific heterogeneous slope showing the effect of a change in the (S/Y) on the (I/Y) and is defined as the savings–retention coefficient. A statistically significant low value of $\beta$ indicates a relatively high degree of capital mobility. A savings–retention coefficient of one implies zero capital mobility as investment is financed by domestic savings. In (1), $\alpha_{1i}$ is a country-specific intercept. Furthermore, the disturbance term $\mu_{i,t}$ in (1) consists of unexplained components of the investment ratio influenced by a set of common factors $f_t$ in (2) with heterogeneous factor loadings of $\tau_i$ that may comprise country-specific fixed effects and heterogeneous country-specific deterministic trends, and the residuals $\varepsilon_{i,\ t}$. In Eq. (2), $\varepsilon_{i,\ t}$ are idiosyncratic disturbance terms distributed with zero mean and finite variances. Furthermore, in Eq. (2), it is highly likely that $f_t$ may induce cross-sectional dependence between the error terms and the explanatory variable $(S/Y)_{i,t}$. Since both the dependent and the explanatory variables are affected by the same unobservable processes $f_t$, the problem of endogeneity may occur in model (1). To avoid this simultaneity problem and the presence of heterogeneity of slope coefficients, the CCEMG estimator adds cross-sectional averages of both the dependent and the explanatory variables to approximate the unobservable factors in running the OLS regression to estimate model (1). Since the CCEMG estimator, unlike a static situation, does not yield consistent estimates of both $\beta_i$ (individual slope estimates) and $\beta_{CCEMG}$ (the average slope coefficient) in a dynamic setting, Chudik and Pesaran (2015) incorporate extra lags ($P_T$ = cube root of T) of the cross-sectional averages of the lagged dependent and explanatory variables as:

$$\left(\frac{I}{Y}\right)_{i,t} = \alpha_i + \lambda_i \left(\frac{I}{Y}\right)_{i,t-1} + \beta_i \left(\frac{S}{Y}\right)_{i,t} + \sum\nolimits_{i=0}^{P_T} \delta'_{i,l} \bar{z}_{t-1} + e_{i,t} \tag{3}$$

where $\beta_i$ represents the DCCEMG estimator, $\beta_{DCCEMG.}$ The estimator, $\beta_{DCCEMG}$ controls for a dynamic panel with lagged dependent variable and weakly exogenous explanatory variables, common effects, and heterogeneity. $P_T$ is the number of incorporated lags, where $\lambda_i$ and $\beta_i$ are stacked into $\pi_i$ = ($\lambda_i$, $\beta_i$). The mean group estimates of DCCEMG are computed as π̂DCCEMG[[]] = 1N ∑i=1 Nπ̂i[[]] (Ditzen 2016, 2018). z¯ [[]] is denoted as:

$$\bar{z}_t = \left[\left(\frac{\bar{I}}{Y}\right)_{t-1}, \left(\frac{\bar{S}}{Y}\right)_{i,t}\right] \tag{4}$$

Although our main objective is to test the presence and degree of capital mobility by applying the DCCEMG estimator, for comparison, we also report the estimates of savings–retention coefficients, $\beta_i$, employing the CCEMG. As shown by Chudik and Pesaran (2015), the CCEMG and DCCEMG approaches offer several econometric advantages in panel data model estimations. Monte Carlo simulations have shown that extending the CCEMG approach to dynamic panels having multifactor error structures performs well, judging by the criteria of biasness, size, power, and root mean square (RMSE), even in samples with low dimensions of N and T (Chudik and Pesaran 2015). Furthermore, as shown by Stock and Watson (2008), the DCCEMG and CCEMG approaches are robust to the presence of structural breaks in the data generating processes (for details, see Eberhardt and Bond 2009; Eberhardt and Teal 2011).

Of note, the CCEMG panel data estimator controls for the existence of cross-sectional dependence that may be found in a panel due to many observed and unobservable common factors, such as aggregate demand and aggregate supply shocks, oil shocks, financial crises, global and local technological shocks, global and local spillover effects, terroristic events, and business cycles. As shown by Kapetanios et al. (2011), the CCEMG estimator, in a static multifactor error structure framework, and the DCCEMG estimator, in a dynamic setting, in addition to handling heterogeneity in slope coefficients and the impact of unobserved common factors, yield consistent estimates regardless of whether the common factors are stationary or non-stationary. Furthermore, the CCEMG and DCCEMG approaches to the estimation of the panel data are robust to the presence of heteroskedasticity, serial correlation, and structural breaks (Pesaran 2006). In the literature, it has been observed that in terms of statistical power, size properties, and RMSE, the CEMG, DCCEMG, and the Gengenbach et al. (2008, 2016) estimators perform well; for example, Chudik and Pesaran (2015), Westerlund and Urbain (2015). Therefore, here, DCCEMG is considered as the preferred panel data estimator.

The data on (I/Y) it, gross capital as a percentage of GDP, and $(S/Y)_{it}$, gross domestic saving as a percentage of GDP, for 20 Latin American and Caribbean countries for 1960 to 2017, are gathered from the World Bank's World Development Indicators (WDIs), 2018 (World Bank 2018). The countries analyzed are Argentina, Bolivia, Brazil, Barbados, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Guatemala, Guyana, Honduras, Jamaica, Mexico, Nicaragua, Paraguay, Peru, Trinidad and Tobago, and Uruguay. In developing countries, there are no available data on investment as a ratio of GDP; instead, we use gross capital as a percentage of GDP. This approach to measure (I/Y) is consistent with previous studies of the FH puzzle in developing countries. In addition to the 2018 WDIs, we use other annual World Bank WDIs as needed.

## Empirical results

Table 1 presents important descriptive statistics on our panel data. Only 4 out of the 20 countries show negative savings in certain years in the period, having spending that depends on borrowing. These are El Salvador, Jamaica, Nicaragua, and Guatemala. Applying the panel-normality test, the Sktest, equivalent to the Jarque-Bera test in panel data, we fail to reject the null hypothesis of normality for the series of *I/Y* and *S/Y*.

**Table 1** Descriptive statistics of the sample

| Variable | Mean | Maximum | Minimum | Std. Deviation |
|---|---|---|---|---|
| (*I/Y*) | 19.44 | 51.75 | 3.39 | 4.91 |
| (*S/Y*) | 17.70 | 59.73 | −9.18 | 8.10 |
| Panel Tests for Normality: Sktest | | | | |
| Joint Test on e: Chi2 (2) = 3.30 *p*-value (0.19) | | | | |
| Joint Test on u: Chi2 (2) = 2.98 *p*-value (0.23) | | | | |

The results of the Pesaran (2007) cross-sectional dependence tests and of the Pesaran and Yamagata slope homogeneity test (Pesaran and Yamagata 2008)[1] reveal cross-sectional dependence and slope heterogeneity in the series. Generally, panel data estimators, such as the fixed effects (FE) and random effects (RE), assume that the slope coefficients are homogeneous. Applying the FE and RE estimators to our data set would lead to biased and inconsistent results. Therefore, we apply the DCCEMG and CCEMG to estimate the average (mean) savings–retention coefficient for both the panel as a whole and the individual countries in the sample.

Furthermore, to ascertain empirically the integration order of variables in levels, we conduct the cross-sectional Pesaran (2007) unit root test (CIPS). The CIPS test is a second-generation panel unit root test that allows for cross-sectional dependence in the data and is robust to the presence of common factors and serial correlation (see Breitung and Pesaran 2008; Pesaran 2007). The CIPS test results are reported in Table 2. The observed *p*-values of the test statistics for various deterministic terms show that both the *I/Y* and *S/Y* series are non-stationary in levels and stationary in first-differences at 1% significance; thus, they are integrated in order of one, I (1).

Once we find that *I/Y* and *S/Y* variables are non-stationary in levels, we test whether there is any long-run economic equilibrium relationship between these variables or whether they are cointegrated through cross-sectional dependence. To that end, we conduct the Westerlund (2007) and Gengenbach et al. (2008, 2016) cointegration tests; the results are reported in Table 3. The Westerlund's cointegration test, in addition to allowing for cross-sectional dependence, is flexible enough to accommodate a large degree of heterogeneity in both the short-run and long-run dynamics (see Westerlund 2007; Persyn and Westerlund 2008). Of note, the Westerlund (2007) cointegration test allows for multiple structural changes in generating data and maintains the null hypothesis of no cointegration. Specifically, it consists of two sets of tests: the group-mean tests ($G_\tau$ and $G_\alpha$) maintaining the alternative hypothesis that at least for one panel member the variables are cointegrated. The other set represents the panel tests ($P_\tau$ and $P_\alpha$), stating the alternative hypothesis that for the panel as a whole *I/Y* and *S/Y* are cointegrated.

The Gengenbach et al. (2008, 2016) test is the panel error correction estimation and has many advantages over Westerlund's (2007) method. This procedure, based on structural dynamics, allows for cross-sectional dependence, non-stationary common factors, and parameter heterogeneity in a multifactor error structure framework for both the individual countries and the panel. In these tests, the maintained null hypothesis is no cointegration (no error-correction) as opposed to the alternative hypothesis of cointegration or the existence of error-correction, for technical details, see Gengenbach et al. (2008, 2016).

The results of the Westerlund (2007) tests are reported in the first panel and the results of the Gengenbach et al. (2008, 2016) tests are reported in the second panel of Table 3. The results of the Westerlund (2007) tests indicate that the observed Z, P, and robust *P*-values are significant at 1% indicating that the null hypothesis of no cointegration is clearly rejected. Thus, we find that the I/Y and S/Y variables do form a long-run

[1]The results of the Pesaran (2007) cross-sectional dependence tests and of Pesaran and Yamagata's slope homogeneity test (2008) are presented in Appendix.

**Table 2** Pesaran's panel unit root CIPS (2007) tests

| Variable | CIPS (No Trend) | CIPS (with Trend) | Determination |
|---|---|---|---|
| $(I/Y)_{it}$ | − 0.761 (0.027) | −2.039 (0.021) | I (1) |
| $(S/Y)_{it}$ | 1.552 (0.940) | 0.590 (0.722) | I (1) |
| $\Delta(I/Y)_{it}$ | −8.199[a](0.000) | −10.421[a](0.000) | I (0) |
| $\Delta(S/Y)_{it}$ | −3.966[a](0.000) | −6.275[a](0.000) | I (0) |

Null hypothesis is the presence of a unit root. Four lags are included
[a]denotes statistical significance at the 1% level

**Table 3** Cointegration tests

| 1. The Westerlund panel cointegration test (2007) results | | | | |
|---|---|---|---|---|
| Test | Value | *Z*-Value | *P*-Value | Robust *P*-Value |
| $G_\tau$ | 2.919 | −3.044 | 0.001 | 0.000 |
| $G_\alpha$ | −15.849 | −2.524 | 0.006 | 0.010 |
| $P_\tau$ | −13.224 | −4.139 | 0.000 | 0.000 |
| $P_\alpha$ | −15.906 | −5.086 | 0.000 | 0.000 |
| 2. Gengenbach et al. (2016) panel and individual country error correction test results | | | | |
| Country | $\alpha_{yi}$ | T $\alpha_{yi}$ [a] | | |
| Argentina | −0.522 | −4.02** | | |
| Bolivia | −0.345 | −3.234 | | |
| Brazil | −0.417 | −3.850*** | | |
| Barbados | −0.283 | −3.08 | | |
| Chile | −0.207 | −2.08 | | |
| Colombia | −0.529 | −4.702* | | |
| Costa Rica | −0.674 | −5.266* | | |
| Dominican Republic | −0.437 | −3.549*** | | |
| Ecuador | −0.201 | −1.919 | | |
| El Salvador | −0.729 | −4.439* | | |
| Gautemala | 0.215 | −2.319 | | |
| Guyana | −0.484 | −4.195** | | |
| Honduras | −0.429 | −3.305 | | |
| Mexico | −0.478 | − 3.624*** | | |
| Jamaica | −0.393 | −4.383** | | |
| Nicaragua | −1.229 | −5.201* | | |
| Paraguay | −0.240 | −2.692 | | |
| Peru | −0.374 | −4.057** | | |
| Trinidad and Tobago | −0.398 | −3.228 | | |
| Uruguay | −0.735 | −5.229* | | |
| **Panel (ECT for Investment)** | −0.466 | −3.719* | | |
| LR Savings Retention Coefficient | 0.240 | 2.54** | | |
| CD | −3.74 (0.000)[b] | RMSE =2.80 | | |

$\alpha_{yi}$ = the error-correction term for the $i_{th}$ country. Null hypothesis is no cointegration. The robust *p*-values are based on 200 bootstrap replications (with lags (1), leads (1), lrwindow (1). The results are robust with other orders of lags, leads and lrwindow and 400 bootstrap replications. $\alpha_{yi}$ denotes the error-correction term
[a] The lag length of 4, based on $[4(T/100)^{2/9}]$ (See Gengenbach et al. 2016). *, ** and *** denote statistical significance at the 1%, 5% and 10% levels of significance, respectively
[b] The observed *p*-value in the parenthesis. Critical values are taken from Tables 1 and 3, with m = 1 and no deterministic term (Gengenbach et al. 2008)

link in the panel. Of note, especially for the Westerlund's group-mean test, the alternative maintains that S/Y and I/Y have a stationary relationship in at least one country in the panel. This implies that in the long run, the solvency condition is satisfied for the panel as a whole (see Coakley et al. 1996; Coakley and Kulasi 1997).

The results of the Gengenbach et al. (2008, 2016) tests illustrate that for the panel as a whole, the observed error correction term rejects the null hypothesis of no cointegration at 1% significance. The observed error correction terms are statistically significant at the 1% level for five panel members and statistically significant at the 5% level for four countries. The long-run savings–retention coefficient we find, reported in Table 3, is much smaller than what has been reported in the literature (see Murthy 2009).

As stated, we use the CCEMG and DCCEMG estimators to measure the degree of capital mobility in the panel and individual countries. Table 4 presents our empirical

**Table 4** The savings retention coefficients

| | | |
|---|---|---|
| 1. Panel estimation | | |
| | CCEMG | DCCEMG[b] |
| Estimated β[a] | 0.21 (4.27)[c] | 0.19 (3.53)[c] |
| Diagnostic Measures | | |
| RMSE | 3.32 | 3.22 |
| $CD_p$ | −4.25 | −4.58 |
| 2. Individual countries estimation | | |
| Country | CCE | DCCE |
| Argentina | −0.65 (0.18) | 0.63 (0.19) |
| Bolivia | 0.74 (1.20) | −0.03(−3.56)[c] |
| Brazil | 0.41 (0.47) | 0.43 (0.50) |
| Barbados | −0.31 (0.47) | −0.28(− 0.52) |
| Chile | 0.38 (0.17) | 0.41 (017) |
| Columbia | 0.52 (0.36) | 0. 57 (0.74) |
| Costa Rica | 0.12 (4.37)[c] | −0.04(−3.93)[c] |
| Dominican Republic | 0.23 0.43) | 0.18 (0.58) |
| Ecuador | 0.21 (0.61) | 0.31 (0.46) |
| El Salvador | 0.09 (0.76) | 0.09 (0.78) |
| Guatemala | 0.04 (4.31)[c] | 0.0.01 (19.98)[c] |
| Guyana | 0..24 (0.21) | 0.27 (0.20) |
| Honduras | 0.21 (0.39) | 0.18 0.52) |
| Jamaica | 0.21 (0.31) | 0.11 (0.81) |
| Mexico | 0.06 (2.25)[c] | −0.12-1.53) |
| Nicaragua | 0.21 (0.42) | 0.22 (0.42) |
| Paraguay | 0.01 (6.23)[c] | −0.07(− 0.98) |
| Peru | 0.0.53 (0.13) | 0.44 (0.18) |
| Trinidad & Tobago | 0.09 (0.54) | 0.13 (0.44) |
| Uruguay | 0.29 (0.38) | 0.34 (0.35) |

[a] statistically different from unity (1) at the 1% level
[b] Recursive mean adjustment and instrument variable adjustment for simultaneity (see, Ditzen 2016). RMSE is the root mean standard error. $CD_P$ is the Pesaran's cross section dependence test statistic (2007). The deterministic terms used in estimations are a constant and a trend. The lag length is 4 (See Chudik and Pesaran 2015)
[c] denotes statistical significance at the 1% level of significance. The observed *Z*-values are in parentheses. CIPS on the residuals using a lag length of 3 for CCEMG = −4.787 and DCCEMG = --6.120, indicate that the residuals are stationary at the 1% level of significance

results of these panel data estimators that take into account common factors, cross-sectional dependence, and slope heterogeneity. The mean savings–retention coefficient of 0.21 and its significant difference from one indicates a relatively high degree of capital mobility in the short run in the panel countries. This finding shows that in the long run, the current account balance is maintained, as the countries in the panel cannot finance their deficits forever. According to Corbin (2004), "...the existence of a one-to-one relationship between the saving and investment rates does not rule out the possibility of lags in the adjustment process of the current account imbalances, which can be viewed as evidence of the existence of capital mobility in the short run" (p. 271).

At the outset, we observe that the magnitudes of the savings–retention coefficients using the CCEMG and DCCEMG estimators are consistently similar around 0.20 and statistically significant, indicating a relatively high degree of capital mobility in the panel countries. The fact that these estimates of the savings–retention coefficient are similar in magnitude and sign indicates that our results are robust to various assumptions. The size and direction of the savings–retention coefficient are reasonable given the economic reforms, improvements in macroeconomic fundamentals, increased financial and economic integration of these countries due to globalization, disinflation, and strong fiscal and monetary policies in effect to create macroeconomic stability. Furthermore, we observe that our estimated savings–retention coefficient, reported in Table 4, is much lower than that reported in the literature. For comparison, Murthy (2009), using the Pedroni panel group mean FM-OLS, reports a savings retention rate of 0.46 with common dummies and 0.48 without common dummies. Cavallo and Pedemonte (2015), employing the FM-OLS estimator on a panel of 24 Latin American and Caribbean countries for the period 1980–2013, report a savings–retention coefficient of 0.39. Payne and Kumazawa (2005), using the panel mean group (PMG) estimator for 19 Latin American countries, report a savings-retention coefficient of 0.35. Our estimate of the savings-retention coefficient, using robust panel econometric procedures, is the lowest among these studies. This is broadly consistent with the financial and global integration that has been taking place in the Latin American and Caribbean countries.

It is also notable that the extent of cross-sectional dependence as measured by the Pesaran CD tests, reported in Table 4, has dropped remarkably, and is barely significant at the 1% level. The presence of a relatively persistent low degree of CD can be attributed to the larger time-series dimension in the data matrix compared with the number of cross-sections. The RMSE associated with all the estimators is relatively low, although it is the lowest for the CCEMG. Furthermore, a low degree of the residual CD may be due to the ever-increasing financial, technological, and economic integration among the countries included in the study. To control for small sample bias, we use a bias correction method, the recursive mean adjustment (see Ditzen 2016; Chudik and Pesaran 2015) in applying the DCCEMG technique for the savings–retention coefficient estimation.

The results of both the CCEMG and DCCEMG estimations of the savings–retention coefficients for individual panel members are reported in the second panel of Table 4. The econometric results for the group as a whole are satisfactory. Savings–retention coefficients for many individual countries are found to be positive and not statistically significant at the lowest 5% level employing both techniques. In panel estimation, what really matters is that the panel (group) coefficient is significant.[2] In conclusion, the estimated $\beta$ coefficients from the Gengenbach et al. (2008, 2016) procedure, the CCEMG, and the DCCEMG estimators are all below 0.21, indicating a relatively high degree of capital mobility.

The main contribution of our study lies in how we test the presence of capital mobility in a panel of 16 Latin American and 4 Caribbean countries by estimating statistically discernible heterogeneous savings–retention coefficients. Using a panel model in a multi-factor error structure consisting of both cross-sectional dependence and unobserved common factors, we apply the DCCEMG estimator. For the degree of capital mobility, the presence of heterogeneity and cross-sectional dependence matter, as stated by Coakley et al. (2004): "It seems that when country heterogeneity, CS dependence, and permanent shocks are explicitly accommodated in a panel framework, the traditional FH puzzle results are completely overturned" (p. 587). Our study attempts to follow Coakley et al.'s (2004) suggestion in testing capital mobility in Latin American and Caribbean countries by applying panel estimation methodologies that allow for slope heterogeneity, cross-sectional dependence, non-stationarity, and endogeneity.

## Conclusion

In this paper, we discuss our investigation of capital mobility in 16 Latin American and 4 Caribbean countries for 1960 to 2017, employing the CCEMG and the DCCMG panel data techniques. These estimators, unlike the widely applied FE and RE methods, are efficient, unbiased, and consistent in the presence of cross-sectional dependence, slope heterogeneity, simultaneity, and non-stationarity. The results of the various tests shared in the paper, point out that our panel suffers from slope heterogeneity, cross-sectional dependence, simultaneity, and non-stationarity. Unlike previous studies on capital mobility in Latin American and Caribbean countries, we show that the magnitude of the savings–retention coefficient is small indicating a relatively high degree of capital mobility, a sign of an integrated capital market. The evidence of cointegration shows that for the countries in the sample, the long-run solvency condition is satisfied. Giannone and Lenza (2008) have demonstrated that when heterogeneous propagation and transmission of global shocks take place, the size of the savings–retention coefficient decreases.

However, it is important to understand that results of the estimated savings retention coefficient illustrate the average across countries and through time-series dimension, indicating a decline in the coefficient for the current period compared with early periods due to changes in financial policies of the Latin American and Caribbean countries. The average of the coefficient illustrates that at the current period it may be even lower than 0.21, taking into account its high level in early periods.

The low savings–retention coefficient reported here implies a higher degree of capital mobility, which is consistent with the changing economic environment in these countries as reflected by economic reforms, structural adjustments, removal of capital controls, and the rapid global development of information technology. Our basic results are further verified to be robust by applying an alternative estimator of panel error correction modeling.

[2]"The standard errors for the individual coefficients are standard OLS standard errors. Even if those are not significant, the mean group estimates can be significant. The reason lies in the different assumptions of the asymptotics of the two coefficients" (from our correspondence with Ditzen, April 2019).

Our results suggest that in recent years, frictions in the financial and goods markets and a degree of "home country bias" in these countries has decreased. In Latin America, since 1990, significant liberalization has been underway (Estevadeordal and Taylor 2013). Some of the prominent liberalization policies include the slashing of tariff and non-tariff barriers, export diversification, deregulation, ambitious preferential trade agreements, such as the 2012 Pacific Alliance agreement between Chile, Colombia, México and Peru, the Central American Free Trade Agreement (CAFTA), the creation of the Custom Union consisting of Argentina, Brazil, Paraguay, Uruguay, Bolivia, and Venezuela, NAFTA (among the United States, Mexico, and Canada), and MERCOSUR (Agreements among Brazil, Argentina, Uruguay, and Paraguay).

These financial integration–augmenting efforts have pushed liberalizing in the movements of goods, services, and capital within the region, in addition to generating externalities creating a conducive environment for increased opportunities for the inflow of foreign domestic investment (FDI). Furthermore, trade agreements stimulate exports, provide legal protection for the properties of enterprises under international law. These agreements also increase competition among firms, engendering cheaper and higher quality goods for consumers, and reduce risks related to a potential escalation of higher tariffs and expropriation of foreign-owned investments. In fact, many economists contend that trade deals and agreements, by reducing frictions that impede trade in goods and services, encourage specialization based on the principle of comparative cost advantage. Ponce (2006), in his econometric study dealing with the impact of foreign trade agreements (FTA) on FDI in Latin American countries during 1985 to 2003, finds the coefficient of FTA highly significant at a 1% level, implying that as countries sign more FTAs, they attract more FDI. The Pacific Alliance, founded in 2011, has been instrumental in increasing its members' global trade from \$876 million in 2010 to \$1.03 trillion in 2016. Such measures have reduced the friction in trading in goods and services, increased productivity, and resulted in the opening up of the economies in Latin American countries. Experts on Latin America, such as Estevadeordal and Taylor (2013), have shown that liberalization has led to the region's median increase in trade GDP ratio by 28%. They also contend that liberalization is estimated to have increased Latin America's GDP per capita growth rate from 0.60 to 0.70 percentage points. In fact, contrary to the Heckscher-Ohlin hypothesis of the substitutability of trade and capital flow, the economic performance of several Latin American countries, especially after the liberalization movement, has shown that capital flow and the consequent FDI are complements rather than substitutes (for detailed effects of liberalization on Latin American countries, see Bown et al. 2017). Therefore, we reason that the econometric finding of a high degree of capital mobility in the Latin American and Caribbean countries during the period investigation is due in part to reduced frictions in goods and services markets leading to increased competition and better access to financial markets, facilitated by regional free trade agreements and other liberalization efforts.

Increased capital mobility in these countries can promote economic growth and hasten the process of globalization by creating a conducive economic environment for FDI in these countries. With increased capital mobility, these countries need not be constrained by a low level of domestic savings based on the low level of income of most citizens. Of course, increased capital mobility, in addition to facilitating more efficient allocation of both physical and financial resources by promoting credit and risk-sharing across international borders, is not without shortcomings. It might induce more economic volatility in the prices of securities, interest rates, and exchange rates. Moreover, the countries cannot simultaneously accomplish their three major economic objectives of having free capital mobility, a fixed exchange rate, and an independent monetary policy. In light of increased capital mobility in these countries, the Central bankers may not pursue monetary policy actions by simultaneously trying to attain their economic external and internal targets such as the exchange rate and interest rate, respectively. However, increased capital mobility would compensate for certain negative outcomes by providing the overall benefits of supplementing insufficient domestic savings and lowering the cost of capital, ultimately leading to increased economic growth and greater financial integration.

Our empirical findings support the prevailing theoretical notion that if the magnitude of the statistically significant savings–retention coefficient is small, then the implication is that in these countries, capital is relatively mobile. The evidence also points to the absence of the FH puzzle. Our study is focused on the estimation of capital mobility in Latin America and Caribbean countries employing recently developed robust panel estimation approaches using the longest period employed in the literature on capital mobility estimations for Latin American and Caribbean countries. However, the employed panel estimation technique does not take into account structural breaks that may exist in developing countries. The potential area for further research is consideration of structural breaks in capital mobility estimations for Latin American and Caribbean countries.

## Appendix

In the first panel of Table 5 the results of the cross-sectional dependence (CD) tests (see, Pesaran 2007) are reported. The observed test statistics of the LM, LM adjusted and Pesaran's CD, for both deterministic terms of trend and without trend in the data generating processes of *I/Y* and *S/Y*, strongly reject the null hypothesis of the presence of cross-section independence at the 1% level significance. The second panel of Table 5 the results of the recently developed Pesaran and Yamagata's slope homogeneity test (Pesaran and Yamagata 2008) are presented. The observed *p*-values of the Δ and the Δ adjusted test statistics clearly indicate that the null hypothesis of slope homogeneity is rejected at the 1% level of significance.

**Table 5** Cross-sectional dependence and homogeneity tests

| 1. Bias-adjusted LM test results for error cross-section dependence[a] | | |
|---|---|---|
| Test | Statistic (No Trend) | With Trend |
| LM | 1101.0[c] | 894.6[c] |
| $LM_{adj}$ | 201.2[c] | 150.0[c] |
| LM CD | 15.01[c] | −13.33[c] |
| 2. The Pesaran and Yamagata (2008) slope homogeneity tests[b] | | |
| Test | Statistic | *P*-value |
| Δ Test | 12.388[c] | 0.000 |
| Δ Adj. Test | 12.715[c] | 0.000 |

[a] Null-hypothesis is Cov ($\mu_{it}$,$\mu_{jt}$) = 0 for all i and i = j
[b] Null-hypothesis is the Homogeneous Slope. Trend is not included
[c] denotes statistical significance at the 1% level

**Author details**
[1]Department of Economics and Finance, Creighton University, Omaha, NE 68178, USA. [2]Department of Economics, Yeditepe University, 26 Agustot Yerlesimi, Atasehir, 34755 Istanbul, Turkey.

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## Publisher's Note