Capital Mobility in Russia

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Abstract

This paper investigates the level of capital mobility in Russia, testing the Feldstein-Horioka puzzle proposed by Feldstein and Horioka (1980). The study examines relations between saving and investment flows in Russia in the presence of structural breaks. It employs the quarterly data for the period 1995-2013, where all estimations are made for two periods, the full period 1995-2013, and 2000-2013, the post Russian crisis period. The empirical analysis includes the Kejriwal and Perron (2008, 2010) structural break test to determine the presence of structural breaks in series and to estimate the saving-retention coefficient under the consideration of structural shifts. To allow for comparison, the parameters of the model were estimated employing the OLS and FMOLS procedures. To test the cointegration relationships between investment and saving flows of Russia, three different cointegration tests were applied to the data. First applied was the Gregory and Hansen (1996) cointegration test, which allows for one structural shift; then, in a case where two breaks were detected, the Hatemi-J (2008) cointegration test was employed. Finally, for a case where more than two breaks are detected, the Maki (2012) cointegration test, which allows for an unknown number of breaks, was applied. The results of this study provide evidence of high capital mobility and reject the existence of the FHP in the post-Russian crisis period. However, failure to reject the hypothesis of no cointegration by all employed cointegration tests denies the solvency of a current account of Russia.

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Key Words: Feldstein-Horioka puzzle, saving-investment association, capital mobility, cointegration, structural breaks, Russia.

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1. Introduction

For the last several decades, the economic crises throughout the world has influenced the rise of global financial integration. Numerous studies have been carried out to investigate capital mobility issues. The most popular concern in capital mobility studies is to explain and solve the Feldstein Horioka puzzle (FHP). Related to the seminal work of Feldstein and Horioka (1980), the FHP established that investment and saving ratios are highly correlated in developed countries and illustrate low capital mobility. These findings are contradict the expected low correlation between investment and savings ratios, particularly in the sample of the OECD developed countries. Since then, a great deal of the attention in the literature has been given to the FHP, with particular focus on European or OECD countries (see, for example, Fouquau et al., 2008; Giannone and Lenza 2008; Kollias et al., 2008; Apergis and Tsoumas, 2009; Kumar and Rao, 2011; Ketenci, 2012, 2013). Apergis and Tsoumas (2009) published the latest updated review of the literature related to the Feldstein Horioka puzzle. The authors conclude that the results of the majority of studies support a high correlation between savings and investments, but at a lower level. At the same time, they indicate that most studies do not validate the capital mobility hypothesis.

For the last several decades, transition and emerging economies have experienced the liberalization process of their economies in trade as well as in capital transactions. However, little attention in the literature has been given to transition and emerging economies, which increasingly are becoming important players in the global financial market (Fidrmuc, 2003; Misztal, 2011; Bose, 2012; Petreska and Mojsoska-Blazevski, 2013). These studies employ panel data obtaining mixed results, while transition and emerging countries are highly heterogeneous. At the same time, they do not include Russia in panel samples. One reasons for this is its large population, compared to the estimated countries, which would significantly affect the average estimations and distort results (Peterska and Mojsoska-Blazevski, 2013). Some authors have included Russia in their comparisons, some of which have been panel studies on the FHP (Aristovnik, 2005; Özmen, 2005; Jamilov, 2013; Trunin and Zubarev, 2013). Therefore, the issue of capital mobility measurements in Russia has not been sufficiently investigated in the literature.

With a population of 143.5 million, Russia is one of the ten most populous countries in the world. In 2012, the GDP value of Russia, was 2.015 trillion USD, which represents 3.25%

of the world economy, putting it on the list of the ten largest world economies.[2] The investigation of capital flows of Russia is not only important on the regional, but on the global level as well. However, there is a lack of studies on capital mobility and its measurement of Russia. Russia is still behind of most advanced countries in terms of free capital mobility; however, it is in front of other emerging countries, such as BRICS,[3] where capital flows are less restrictive (see, for example, Figure 1).

Since the transition started, the capital liberalization policy for capital account has been cautious and gradual in transition countries, where non-FDI related transactions have been restricted. However, Russia has had a different program for capital liberalization compared to that of the Commonwealth of Independent States (CIS), which started the process of transition at the same time. The liberalization of FDI transactions has had strict limitation with gradual ease. Restrictions on non-resident portfolio investments gradually were removed by early 1998. However, during the crisis, some capital restrictions were returned with further gradual liberalization after 2000. Comparing Russia to the CIS, at the beginning of the transition, most of the total net capital flows in the CIS where on account of Russia with continuous increase until the August of 1998 crisis with gradual recovery after 1999.

In terms of structure, foreign direct investments accounted for a small share of Russian capital inflows, while the net short term external liabilities significantly increased before the crisis, followed by decline during the Russian crisis (Buiter, 2003).

Following the gradual liberalization after the crisis, investments grew again. Particularly, capital flows increased sharply after 2004, when the new foreign exchange law came into force, which was directed on the progressive liberalization of capital movements. The new law still had various restrictive capital control arrangements, but they were phased out in 2006 (OECD 2006). Thus, particularly for the period 2004-2008, Russia experienced net capital inflow, where, for example, about a quarter of inward FDI belonged to capital inflows from Cyprus accounts owned by Russian nationals (Brockmeijer et al., 2012). In general, Russia experiences considerable capital outflow of domestic savings to foreign commercial banks; however, despite this high rate of capital outflow, particularly the outflow of domestic savings, in 2013, Russia was ranked the third most attractive country for foreign investors after the US and China, after having been ninth on this list in 2012.[4] The level of

[2] World Bank
[3] BRICS – Brazil, Russia, India, China and South Africa.
[4] UNCTAD, Global Investment Trends Monitor.

capital mobility has increased continuously in Russia; therefore, it is expected that the correlation between investments and domestic savings is low.

The purpose of this article is to make a contribution to the literature on the capital mobility analysis in Russia. The study examines the FHP, employing the latest econometric techniques that accommodate structural breaks. Quarterly data are used, covering the period from 1995 to the third quarter of 2013. Estimates are made for two periods: from 1995 to 2013, is the full period; and from 2000 to 2013, the period during which gradual capital mobility liberalization was applied, or the post-Russian crisis period. The remainder of the paper consists of the following sections: Section 2 outlines the empirical methodology adopted in the paper. Section 3 presents the empirical results, and section 4 draws conclusions on the data.

2. Methodology

This study examines the degree of capital mobility in Russia in the presence of structural breaks. Feldstein and Horioka (1980) first who investigated the level of capital mobility in OECD countries by estimating the following equation:

$$\left(\frac{I}{Y}\right)_i = \alpha + \beta\left(\frac{S}{Y}\right)_i + e_i \tag{1}$$

Where I is gross domestic investment, S is gross domestic savings and Y is gross domestic product of considered country i. Coefficient β which is known as saving retention coefficient measures the degree of capital mobility. If a country possesses perfect international capital mobility, the value of β has to be close to 0. If value of β is close to 1, it would indicate the capital immobility of the country. The results of Feldstein and Horioka (1980) showed that the value of β for 21 open OECD economies changes between 0.871 and 0.909, illustrating by this international capital immobility in considered countries. These controversial results gave start to widespread debates in the economic literature. Numerous studies have provided evidence supporting these results, at the same time different results exist in the literature with a wide array of interpretations. Therefore, the findings of Feldstein and Horioka (1980), which are contrary to economic theory, started to be referred to as “the mother of all puzzles” (Obstfeld and Rogoff, 2000, p.9).

In the long run macroeconomic series including investment and savings may contain a variety of structural changes within a country or at the international level. Therefore in order to examine the regression model (1) in the presence of multiple structural breaks, Kejriwal and Perron (2008, 2010) approach was employed in this study. Kejriwal and Perron (2008, 2010) developed estimation of cointegrated regression models accounting for multiple structural changes. The framework of this approach is general enough to allow for both stationary and non-stationary variables in the model, at the same time it allows for serial correlation and heteroskedasticity. Authors illustrated that inference is possible in models with both stationary and non-stationary variables, as long as the intercept is allowed to change through regimes. Their work is based on Bai and Perron (1998) methodology that estimates and tests linear models of stationary variables for multiple structural changes. Kejriwal and Perron (2008,2010) derived limiting distributions of sup-Wald test of Bai and Perron (1998) under general conditions for errors and regressors to allow for non-stationary variables in cointegrated regressions.

The methodology considers the multiple linear regression in the presence of *m* breaks, which means *m+1* regimes.

$$y_t = x_t'\beta + z_t'\delta_j + e_t \qquad (2)$$

where $t = T_{j-1} + 1, \ldots, T_j$ is the time period with $j = 1, \ldots, m+1$ regimes. y_t is dependent variable of the regression, x_t and z_t are vectors of covariates with sizes of $(px1)$ and $(qx1)$, respectively, β and δ_j are vectors of coefficients, where the parameter vector β is not subject to change, while δ_j is changing across regimes. Finally, e_t is the disturbance term of the regression. The purpose of this methodology is to estimate the unknown coefficients of the regression together with treated as unknown *m* number of break points. For every *m* partition $(T_1, \ldots, T_m)$, estimates of coefficients β and δ_j are generated by minimizing the sum of squared residuals which is represented by the following equation:

$$S_T(T_1, \ldots, T_m) = \sum_{i=1}^{m+1} \sum_{t=T_{i-1}+1}^{T_i} \left[y_t - x_t'\beta - z_t'\delta_t \right]^2 \qquad (3)$$

Substituting estimates $\hat{\beta}(\{T_j\})$ and $\hat{\delta}(\{T_j\})$ into equation (3) the estimators of break locations will be obtained, which are the global minimum of the sum of squared residuals objective function, and can be expressed by the following equation:

$$(\hat{T}_1, \ldots, \hat{T}_m) = \arg\min_{T_1, \ldots, T_m} S_T(T_1, \ldots, T_m) \qquad (4)$$

The minimization of the sum of squared residuals is obtained in all partitions $(T_1, \ldots, T_m)$, that $T_i - T_{i-1} \geq q$. The estimates of regression parameters are least-squares estimates associated with *m*-partition $\{\hat{T}_j\}$, i.e. $\hat{\beta} = \hat{\beta}(\{T_j\})$ and $\hat{\delta} = \hat{\delta}(\{T_j\})$. Bai and Perron

(2003) proposed the efficient algorithm of obtaining the locations of break points, which is based on the principle of dynamic programming.

The procedure for the specification of the number of breaks proposed by Bai and Perron (1998) is as follows. Firstly, the statistics for *UD*max and *WD*max tests have to be calculated. *UD*max and *WD*max tests are double maximum tests that examine for the hypothesis of no structural break against an unknown number of breaks with the given upper bound of breaks *M*, and can be calculated by the following formulas:

$$UD\max F_T(M,q) = \max_{1\le m\le M} \sup_{(\lambda_1,\dots,\lambda_m)\in\Lambda_\varepsilon} F_T(\lambda_1,\dots,\lambda_m;q). \tag{5}$$

where $F_T(\lambda_1,\dots,\lambda_m;q)$ is the sum of *m* dependent chi-square random variables, each one divided by *m*, with *q* as degree of freedom.

$$WD\max F_T(M,q) = \max_{1\le m\le M} \frac{c(q,\alpha,1)}{c(q,\alpha,m)} x \sup_{(\lambda_1,\dots,\lambda_m)\in\Delta_\varepsilon} F_T(\lambda_1,\dots,\lambda_m;q). \tag{6}$$

where *c(q, α, m)* is the asymptotic critical value of the individual tests with *α* as significance level.

Next, Wald type tests have to be applied, where the sup F(0|1) test examines for the hypothesis of no breaks against 1 break existence. If the statistics of this test reject the hypothesis of no breaks, the sup F(*l+1*| *l*) has to be applied to specify the number of breaks in series. The number of breaks in series can be chosen as well on the basis of the Bayesian Information Criteria (BIC), and the modified version of BIC proposed by Liu et al. (1997) (LWZ).

Before proceeding to cointegration tests, the stationarity of employed variables has to be examined. In order to test integration properties of variables two different unit root tests were applied. The first test is the unit root test proposed by Ng and Perron (2001), which has maximum power against *I*(0) alternatives. In order to generate efficient versions of the modified tests of Perron and Ng (1996), Ng and Perron (2001) employed the generalized lest squares detrending procedure proposed by Elliot, Rothenberg and Stock (1996). Ng and Peron stressed that the choice of the lag length of a regression is extremely important for the good size and power properties of an efficient unit root test. Therefore, Ng and Perron proposed modified AIC and recommended the use of a minimized value of modified Akaike information criterion (AIC) for selecting the regression's lag length.

An additional unit root test employed in this study is a test proposed by Zivot and Andrews (1992), which is the sequential break point selection test with the null hypothesis of unit root without structural break against the alternative that series are trend-stationary with one break point. Zivot and Andrews considered three different models: model A allows for a break in the intercept; model B allows for a break in the slope; and model C allows for a single break in the intercept and in the slope of the function. In this study, model C was employed.

Cointegration

Finally, in order to test for cointegration characteristics between variables under the consideration of a structural break presence, the Gregory and Hansen (1996) test was employed for a case where one structural shift was detected. This test allows for the break in the three alternative models, such as a break in the level (model C), in the level with trend (model C/T), and in the level and slope coefficients (model C/S). For a case where the Bai and Perron (1998) test detected two breaks, the Hatemi-J (2008) test was employed. The Hatemi-J (2008) test is an extended procedure of the Gregory and Hansen (1998) method to allow for two structural shifts in three different models: model C, model C/T and model C/S.

For a case where more than two breaks were detected, the Maki (2012) test was applied. The Maki (2012) test is based on the Bai and Parron (1998) test for structural breaks, and on the unit structural breaks proposed by Kapetanios (2005). The Maki (2012) proposes cointegration tests allowing for an unknown number of breaks. The null hypothesis of the test is no cointegration, with the alternative hypothesis of cointegration with unspecified number of breaks i that are smaller or equal to the maximum number of breaks ($i \leq k$). The Maki (2012) test has an advantage over standard cointegration tests that allow for one or two structural changes in cointegration relationships when multiple unknown numbers of breaks exist.

3. Empirical Results

Unit root tests

Table 1 presents the results of the Ng and Perron (2001) unit root tests. Results are presented for two considered periods, 1995-2013 and 2000-2013. All tests are consistent with each other and the null hypothesis of the unit root presence was not rejected by any of the tests for any of the employed variables, investments and savings, and for any of considered periods. Next, the Zivot and Andrews (1992) unit root test, which allows for a structural break allocation, was applied to series for both periods. The t statistics of the test and possible break allocation are presented in Table 2. When a structural break is allowed, the unit root hypothesis also was not rejected for considered series and periods. The results of the unit root tests demonstrate the non-stationarity of the employed variables in both periods. Having verified the non-stationarity of the series under observation by the Ng and Perron (2001) and the Zivot and Andrews (1992) unit root tests, structural change presence and cointegration tests were conducted.

Structural change presence

The Kejriwal and Perron (2008, 2010) methodology allows for the presence of non-stationary as well as stationary variables; however, it was developed for cointegrated regression models. Therefore, before proceeding to the structural change presence test, first, it is important to estimate the cointegrating relationships of variables. For this reason, the Johansen cointegration test was conducted. In order to determine the rank of cointegration space, two test statistics are presented, the Trace and the Max-Eigenvalue (Table 3). The results of the Trace likelihood ratio test statistic and of the Max-Eigenvalue likelihood ratio test statistic were consistent with each other. The results of the tests indicated two cointegration relationships at the 5% significance level between saving and investment variables for the 1995-2013 period. For the second period, 2000-2013, the estimation results revealed one cointegrating equation at the 5% significance level and two cointegration equations at the 10% significance level. Thus, the results of Table 3 indicate the existence of long-run relationships between chosen variables in all cases when structural breaks are not taken into account.

Having verified the existence of long-run relationships between the variables, the Kejriwal and Perron (2008, 2010) methodology was applied to the series. Table 4 reports the results of the Kejriwal and Perron (2008, 2010) tests for detecting structural changes. Sup F(k) tests are significant for all values of k in both periods, except when k =1 in the second considered period. The last two columns of the table present statistics for the *UD*max and *WD*max tests that are significant in both periods as well. Once more, the null of no structural breaks was rejected by both tests. Combining the results of tests presented in Table 4, it can be concluded that there is strong evidence of a structural change presence in the employed series in both considered periods.

Table 5 reports the results for the sequential test l versus $l+1$ structural changes proposed by Bai and Perron (1998). In this study, the sequential test (S), the Bayesian information criterion (BIC), and the modified Schwarz criterion (LWZ) were used for the detection of the number of breaks in series, and their results are presented in last three columns of the table. In the full considered period 1995-2013, the sequential test detected two and the BIC and LWZ detected three structural shifts. In the after crisis period, 2000-2013, the sequential test did not detect any structural shift, while the BIC and LWZ detected one break. Because the Kejriwal and Perron test (Table 4) provided evidence of a structural shift presence, the results of the BIC and LWZ for one structural shift were considered in this study for the 2000-2013 period.

Cointegration

Tables 6 and 7 present the estimation results of the cointegration tests for the 1995-2013 period. The Bai and Perron (1998) test detected two structural shifts with sequential procedure and three structural shifts with BIC and LWZ procedures for the 1995-2013 period. Therefore, first the Hatemi-J (2008) test, which allows for two structural shifts, was applied, and then the Maki (2012) test, which designs cointegration relationships when multiple unknown numbers of breaks are allowed. The results of the tests are presented in Tables 6 and 7, respectively. All of the test statistics of the Hatemi-J test for the three employed models, C, C/T, and C/S, failed to reject the hypothesis of no cointegration, Table 6. The results of the Maki (2012) test are demonstrated in Table 7, where MB*k* presents the t-statistics of the Maki test, where *k* denotes the maximum number of breaks. The test statistics failed to reject the null of no cointegration as well.

The Bai and Perron (1998) test detected one structural shift, with the BIC and LWZ procedures, for the 2000-2013 period. Therefore, the Gregory and Hansen (1996) test was applied to the series of the 2000-2013 period, Table 8. The results of the cointegration test statistics do not provide evidence of cointegration in series in any of considered models, C, C/T and C/S, when a structural break is allowed in the post crisis period.

The results of the cointegration estimations that allow for structural shifts did not provide evidence for the existence of cointegration relationships in any considered period. In the literature, the cointegration presence between savings and investment is interpreted as the long-run solvency condition, which exists regardless of the level of capital mobility, implying the effective realization of government policies targeting a sustainable current account (Coakley et al., 1996; De Vita and Abbott, 2002; Abbott and De Vita, 2003; Vasudeva Murthy, 2009). The disappearance of long-run relationships with the introduction of structural breaks denies the solvency of a current account in Russia in both considered periods.

Coefficients estimates

Table 9 reports the results of the parameters estimations of regression (2) in the presence of structural breaks, where dependent variable y_t is the ratio of gross domestic investments to the gross domestic product, and covariate x_t is the ratio of gross domestic savings to the gross domestic product. Estimates of break locations are given in the last three columns $\{\hat{T}_j\}$ of the table, based on a 95% confidential level. Estimates of the saving retention coefficient, $\hat{\beta}$, corrected for the presence of structural breaks, are given in the second column. In the full estimated period, 1995-2013, the saving retention coefficient was found at a low level, close to zero, or -0.01 when three breaks are detected by the BIC and the LWZ procedures, and 0.05 when two breaks are detected by the sequential test. However, in both cases, the saving retention coefficient estimates were not found significant. Estimations of the post-crisis period 2000-2013 produced significant results for the saving retention coefficient when one structural break was detected by the BIC and LWZ procedures. Thus, the estimate of the saving retention coefficient in the presence of a structural break was found at the level -0.10, which is relatively close to zero.

For comparison, the saving retention coefficient is estimated using the OLS and FMOLS procedures, Table 10. The OLS and FMOLS estimation results are similar and consistent to coefficient estimations with structural shifts allowance. The saving retention coefficient was found not significant in the full considered period, 1995-2013. However, the estimations for the post crisis period revealed significant saving retention coefficient with negative sign at the -0.275, and at the -0.306 levels by the OLS and FMOLS procedures, respectively. The negative sign of the saving retention coefficient can be interpreted as low correlation between saving and investment flows, or as the existence of high saving flight abroad due to domestic financial structure deficiency (Özmen, 2004).

The problem of capital flight in Russia has been present since the early 1990s. Three different examples of domestic capital flight exist: to transfer assets abroad that are denominated in a foreign currency, to accumulate profits from financial assets that are located abroad and denominated in foreign currency, and to transfer financial assets in national currency into financial assets denominated in foreign currency. Domestic capital flight has existed since the Russian economy moved to the market economy model. However, capital flight from Russia mainly is not connected to the normal decision of profit maximization, but rather it can be explained by motivations driven by general or currency risk that lead to significant reduction in national investments (Abalkin and Whalley, 1999).

Except for the period 2004-2008, when Russia experienced net capital inflow and about a quarter of inward FDI were on account of capital inflows from Cyprus accounts owned by Russian nationals (Brockmeijer et al., 2012), capital flight in Russia continues to increase. The net capital outflow for several previous years composed four percent of the GDP can be explained by an unfavourable investment climate. Capital flight from Russia, $133.7 billion in 2008, decreased to $56.1 billion and to $34.4 billion in 2009 and 2010, respectively, and then rose to $80.5 billion and $56.8 billion in 2011 and 2012, respectively.[5] The main concern of domestic capital outflow in Russia is its effect on domestic investments; therefore, in order to cover the gap of the deficit of domestic savings, Russia attempts to attract foreign capital. Thus, in 2013, after the US and China, Russia was accepted as the third most attractive country for foreign investors after having been ninth in this list in 2012.[6] As a result, the level of capital mobility has continuously increased in Russia, decreasing the level of correlation between investments and domestic savings.

The results of the saving retention coefficients estimates illustrate a high mobility of capital in Russia in the post-crisis period. Consideration of structural shifts does not significantly affect estimation results where structural shifts are not allowed. Nevertheless, the allocation of structural breaks in the model may correct estimated parameters for the provision of better capital mobility illustration. Thus, the results of the regression estimates provide rather weak evidence of FHP presence in Russia in the post-crisis period.

The limited literature on the measurement of capital mobility in Russia provides mixed results. For example, Jamilov (2013) estimated the capital mobility of the Caucasus region for the period 1996-2010 employing panel econometric techniques such as the Fully Modified OLS (FMOLS), the Dynamic OLS (DOLS), and the Pooled Mean Group (PMG). However, each panel cointegration estimation method provided different results for the individual countries. Thus, the saving-retention coefficient for Russia was found significant in all three cases, but values were found at different levels, -0.21, -0.02, and 1.49, respectively, to an employed method. Therefore, it is difficult to make a certain conclusion without choosing a particular method. Trunin (2013) investigated the level of capital mobility and the global financial effect for developed and developing countries for the periods 1996-2011 and 2007-2011. The saving retention coefficient for the period 1996-2011 for Russia was not found

[5] Sergei Ignatyev, Chief of the Central Bank of Russia. 05.06.2013 RIA Novosti.

[6] UNCTAD, Global Investment Trends Monitor.

significant at the level 0.221, which is compatible with the present study results. In the post-crisis period 2007-2011, the saving-retention coefficient was found significant at the level 0.8, indicating a low capital mobility level after the global crisis. However, the latest estimations were made employing only five observations, therefore it is not enough to make any certain conclusions about the capital mobility level in this period.

Thus, the results of this study employing OLS and FMOLS estimations provide weak evidence of FHP presence in Russia in the post-crisis period, while estimations with accommodation for structural breaks illustrate high capital mobility and no existence of FHP in the Russian post-crisis period.

4. Conclusion

This paper examined capital mobility in Russia in the presence of structural breaks for two periods, from 1995-2013, and the post-crisis period from 2000-2013. Recently developed econometric methods were applied to quarterly series in order to investigate the cointegrating relationships of investment and savings variables, taking into account the presence of structural shifts in the model when it was relevant and to estimate the saving retention coefficient. The long-run macroeconomic series including investment and saving flows may contain a variety of structural changes within a country or at the international level. Therefore, in order to examine the regression model (1) in the presence of multiple structural breaks, the approach of Kejriwal and Perron (2008, 2010) was employed. Kejriwal and Perron (2008, 2010) developed the estimation of cointegrated regression models accounting for multiple structural changes. The test provided strong evidence of structural shifts presence in employed series in both of the considered periods. Thus, in the period 1995-2013, two shifts were detected by the sequential test, and three shifts by the BIC and LWZ procedures. In the post-crisis period, 2000-2013, one shift was detected by the BIC and LWZ procedures.

To examine the cointegration relationships of series in the presence of structural breaks, the Hatemi-J (2008) and the Maki (2012) cointegration tests were employed. These allow for two and multiple structural breaks, respectively, for the period from 1995-2013. For the post-crisis period 2000-2013, the Gregory and Hansen (1996) cointegration test was employed that allows for the presence of one structural shift. The results of the cointegration estimations that allow for structural shifts did not provide evidence of the existence of cointegration relationships in any considered period. The disappearance of long-run relationships with the introduction of structural breaks denies the solvency of a current account in Russia in both of the considered periods.

The OLS and FMOLS estimates of the saving-retention coefficient and the coefficient estimates of the Kejriwal and Perron (2008, 2010) procedure that are corrected for the presence of structural breaks were not found significant in the full estimated period, 1995-2013. However, estimations of the post-crisis period were found significant with negative sign at the -0.275 and at the -0.306 levels by the OLS and FMOLS procedures, respectively, and at the level -0.10, when a structural break was allowed.

The results of the study indicate the presence of high capital mobility in Russia in the post-crisis period. The negative sign of the saving retention coefficient confirms the high level of domestic capital flight. The consideration of structural shifts does not significantly affect the estimation results where structural shifts are not allowed. Nevertheless, the allocation of structural breaks in the model corrects estimated parameters for the provision of better capital mobility illustration. Thus, the results of this study employing OLS and FMOLS estimations provide weak evidence of FHP presence in Russia in the post-crisis period, while estimations with accommodation of structural breaks illustrate high capital mobility and no existence of FHP in the Russian post-crisis period.

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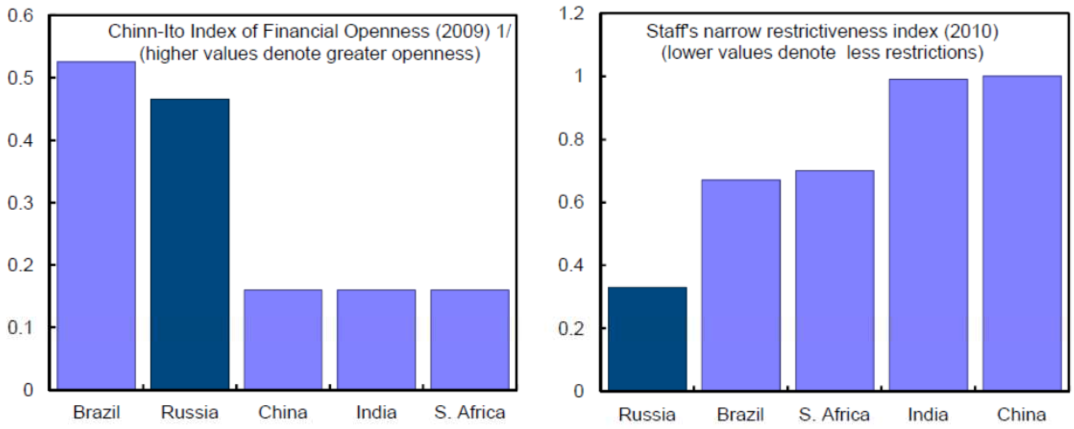
6. Appendix

Figure 1. Russia BRICS – De Jure Capital Flow Restrictiveness



Source: Brockmeijer et al., 2012, Figure 8.

Table 1. Unit Root Tests Ng and Perron (2001)

Period	MZ_α^{GLS}	MZ_t^{GLS}	MSB^{GLS}	MP_T^{GLS}	MZ_α^{GLS}	MZ_t^{GLS}	MSB^{GLS}	MP_T^{GLS}
1995-2013	***Investments***				***Savings***			
Level	-6.54	1.81	0.28	13.93	-13.71	-2.60	0.18	6.74
2000-2013								
Level	-7.68	-1.84	0.24	12.15	-9.59	-2.19	0.23	9.51

Notes: MZ_α^{GLS} is the modified Phillip-Perron test MZ_α; MZ_t^{GLS} is the modified Phillip-Perron MZ_t test; MSB^{GLS} is the modified Sargan-Bhargava test; MP_T^{GLS} is the modified point optimal test. For details, see Ng and Perron (2001). The order of lag to compute the test was chosen using the modified AIC (MAIC) suggested by Ng and Perron (2001). The critical values for the above tests were taken from Ng and Perron (2001).

Table 2. Unit Root Tests Zivot and Andrews

Period	*t* statistics	break	*t* statistics	break
1995-2013	**Investment**		**Savings**	
	-3.513	2000:1	-4.754	1999:1
2000-2013				
	-4.667	2006:4	-4.473	2008:4

Notes: The critical values for the Zivot and Andrews test are -5.57, -5.08 and -4.82 at 1%, 5%, and 10% levels of significance, respectively.

Table 3. Standard Cointegration Test Johansen

Period	**Trace statistics**		**Max-Eigen Statistics**	
	r = 0	r ≤ 1	r = 0	r ≤ 1
1995-2013	27.56**	6.28**	21.28**	6.28**
2000-2013	26.83**	3.43*	23.41**	3.43*

Notes: ** and * denote statistical significance at 5% and 10% levels.

Table 4. Structural Break Tests of Kejriwal and Perron (2008, 2010).

Period	Sup F(1)	Sup F(2)	Sup F(3)	Sup F(4)	Sup F(5)	UDmax	WDmax
1995-2013	38.34**	45.91**	123.47**	76.76**	93.34**	123.47**	204.83**
2000-2013	2.51	18.26**	243.63**	213.22**	33.94**	243.63**	366.62**

Notes: ** denotes statistical significance at 5% level. The 5% critical values for the supF(l) test in the case of non-stationary variables are 14.30, 12.11, 10.41, 9.19 and 7.64 for l = 1,2,3,4,5, respectively. The ritical value for the UDmax test is 14.47. See Kejriwal and Perron (2010). The critical value for the WDmax test is 9.039. See Bai and Perron (2003-1). The 5% critical values for the supF(l) test in the case where stationary and non-stationary variables are allowed are 14.53, 11.94, 10.38, 9.28 and 7.51 for l = 1,2,3,4,5, respectively. The critical value for UDmax test is 14.79.

Table 5. Sequential Test of *l* versus *l+1* Structural Changes.

Period	Sup F(2|1)	Sup F(3|2)	Sup F(4|3)	Sup F(5|4)	S	BIC	LWZ
1995-2013	73.89**	3.75	1.53	0.02	2	3	3
2000-2013	0.13	0.0001	0.0001	0.0003	0	1	1

Notes: ** denotes statistical significance at 5% level. * denotes statistical significance at 10% level.
S - sequential procedure, BIC-Bayesian Information Criteria, LWZ, the modified version of BIC proposed by Liu et al. (1997), are used for the selection of breaks number. The 5% critical values for the sup F($l+1| l$) test are 10.13, 11.14, 11.83 and 12.25 for l = 1,2,3,4 respectively, see Bai and Perron (2003-1) Table 2c.

Table 6. Cointegration Test with Two Structural Breaks Hatemi-J (2008) for the 1995-2013 Period.

Model	ADF*	Z_t^*	Z_α^*
C	**-4.33**	**-4.51**	**-31.17**
1	1999:Q2	1997:Q4	1997:Q3
2	2003:Q3	2004:Q1	2004:Q2
C/T	**-4.71**	**-4.96**	**-36.73**
1	2003:Q1	1997:Q4	1997:Q4
2	2005:Q1		
C/S	**-4.99**	**-4.44**	**-31.33**
1	1999:Q2	1999:Q1	1999:Q1
2	2004:Q1	2004:Q1	2004:Q2

Notes: The critical values are collected from Hatemi-J (2008) are -6.503, -6.015 and -5.653 (1%, 5% and 10%) for the ADF and Z_t' tests, and are -90.794, 76.003 and 52.232 (1%, 5% and 10%) for the Z_α' tests.

Table 7. The Cointegration Test Maki (2012) with Unknown Number of Breaks for the 1995-2013 Period.

MB1	MB2	MB3	MB4	MB5
-4.20	-4.48	-4.59	-4.77	-4.77

Notes: Critical values are taken from Maki (2012) – Table 1.

Table 8. Cointegration Test with a Structural Break Gregory and Hansen for the 2000-2013 Period.

Model	ADF*	Z_t^*	Z_α^*
C	-4.92	4.68	-28.89
C/T	-4.83	-4.71	-29.07
C/S	-4.63	-4.79	-32.67

Notes: Critical values for the Gregory and Hansen test are reported in the Table 1 of Gregory and Hansen (1996).

Table 9. Estimated Regression Parameters under Breaks.

Period	$\hat{\beta}$	$\hat{\delta}_1$	$\hat{\delta}_2$	$\hat{\delta}_3$	$\hat{\delta}_4$	$\hat{T}_1$	$\hat{T}_2$	$\hat{T}_3$
1995-2013								
(BIC, LWZ)	-0.01 (0.03)	20.13** (0.87)	15.85** (0.93)	18.36** (1.01)	21.28** (0.95)	1997:Q3 ('96:Q4-'98:Q1)	2000:Q2 ('99:Q3-'02:Q2)	2006:Q3 ('06:Q1-'07:Q1)
(S)	0.05 (0.04)	18.32** (1.08)	15.47** (1.21)	19.24** (1.17)	-	1997:Q3 ('95:Q3-'97:Q4)	2006:Q3 ('06:Q2-'11:Q3)	-
2000-2013								
(BIC, LWZ)	-0.10** (0.03)	21.17** (1.12)	24.02** (1.03)			2006:Q3 ('05:Q3-'07:Q1)		

Notes: The parentheses under the break points are 95% confidence intervals for the break dates.
**, * Denote statistical significance at the 1 and 5% levels, respectively.
S - sequential procedure, BIC-Bayesian Information Criteria, LWZ - the modified version of BIC proposed by Liu et al. (1997).

Table 10. Estimated Regression Parameters OLS and FMOLS.

Period	*OLS*		*FMOLS*	
	α	β	α	β
1995-2013	20.091** (1.876)	-0.036 (0.060)	19.768*** (3.243)	-0.026 (0.104)
2000-2013	28.251** (2.007)	-0.275** (0.062)	29.261*** (3.784)	-0.306*** (0.117)

Notes: *** and ** denote statistical significance at 1 and 5% levels, respectively.
α and β coefficients are from equation 1.