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Capital mobility in Russia

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Abstract

This paper investigates the level of capital mobility in Russia, testing the Feldstein–Horioka (1980) puzzle (FHP). 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, in which 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 estimate the savings retention coefficient under the consideration of structural shifts. To facilitate comparison, the parameters of the model were estimated employing the OLS and FMOLS procedures. To test the cointegration relationships between investment and saving flows in Russia, two different cointegration tests were applied to the data. The first applied was the Maki (2012) cointegration test, which allows for an unknown number of breaks; then, in a case where only one break was detected, the Carrion-i-Silvestre and Sanso (2006) cointegration test was employed. The results of this study provide evidence of high capital mobility and reject the existence of the FHP in the post-Russian crisis period. Evidence of the cointegration presence indicates the solvency of a current account in Russia.
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1. Introduction

For the last several decades, economic crises throughout the world have been influenced by the rise of global financial integration. Numerous studies have been carried out to investigate capital mobility issues. The most popular concern

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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 savings ratios are highly correlated in developed countries and demonstrate low capital mobility. These findings 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 FHP. The authors conclude that the results of the majority of studies support a high correlation between savings and investments but at a lower level. Meanwhile, 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 in trade and capital transactions. However, little attention has been given in the literature 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, whereas transition and emerging countries are highly heterogeneous. Moreover, they do not include Russia in panel samples. One reason for this is its large population compared with the estimated countries, which would significantly affect the average estimations and distort the 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). However, 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 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.[1] The investigation of capital flows of Russia is not only important at the regional level but on the global level as well. However, there is a lack of studies on capital mobility and its measurement in Russia. Russia is still behind most advanced countries in terms of free capital mobility; however, it is in front of other emerging countries, such as BRICS[2] (see, for example, Fig. 1).

Since the transition began, the capital liberalization policy for capital accounts 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 been executed under strict limitation with gradual ease. Restrictions on nonresident portfolio investments were gradually removed

[1] World Bank.
[2] BRICS — Brazil, Russia, India, China and South Africa.

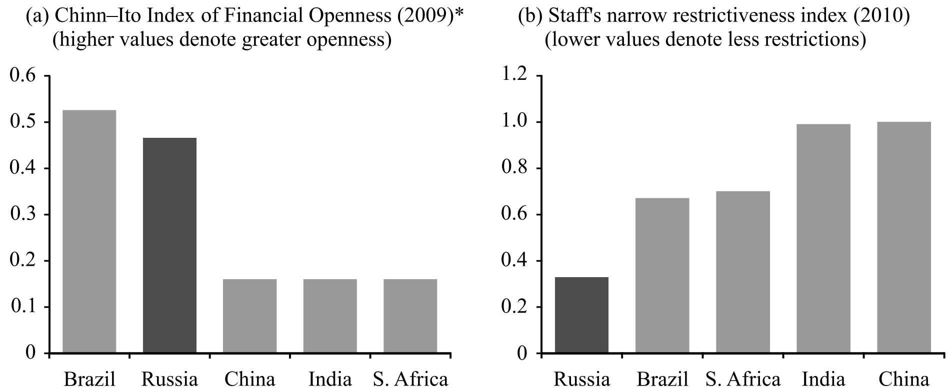

Fig. 1. Russia: BRICS — de jure capital flow restrictiveness.

Note: * Maximum index value is normalized at one.
Source: Brockmeijer et al., 2012, p. 35, Fig. 8.

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 total net capital flows in the CIS involved Russia, with a continuous increase until the August 1998 crisis and gradual recovery after 1999.

In terms of structure, foreign direct investments accounted for a small share of Russian capital inflows, whereas the net short-term external liabilities significantly increased before the crisis, followed by a 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 toward 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, in which, for example, approximately one-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 behind the US and China, after having been ninth on this list in 2012.[3] The level of 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 taken from the Organisation for Economic Co-operation and Development (OECD), Quarterly National Accounts Dataset, covering the period from 1995 to the third quarter of 2013. Estimates are made for two periods: 1995

[3] UNCTAD, Global Investment Trends Monitor.

to 2013 is the full period; and 2000 to 2013 is 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.

2. Methodology

This study examines the degree of capital mobility in Russia in the presence of structural breaks. Feldstein and Horioka (1980) first 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 the gross domestic investment, S is the gross domestic savings, and Y is the gross domestic product of considered country i. Coefficient β, which is known as the saving retention coefficient, measures the degree of capital mobility. If a country possesses perfect international capital mobility, the value of β must be close to 0. If β is close to 1, it would indicate capital immobility within 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 and illustrated the international capital immobility in the considered countries. These controversial results sparked widespread debates in the economic literature. Numerous studies have provided evidence supporting these results, and 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. For example, Fig. 2 illustrates gross domestic investment and gross do-

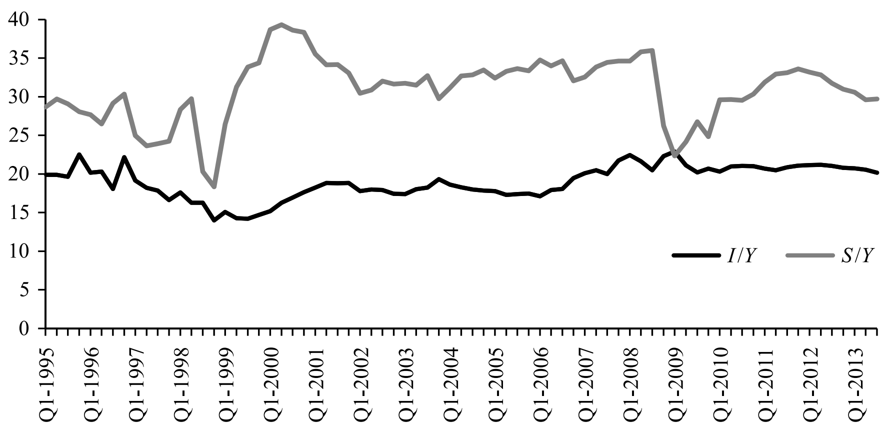

Fig. 2. Gross domestic investment and gross domestic savings in Russia.

Source: Author’s representation of the employed dataset.

mestic savings as a share of GDP in Russia for the period 1995–2013. The graph shows that variables are not correlated — the correlation coefficient is –0.07 — and that variables demonstrate the existence of structural shifts. Therefore, to examine the regression model (1) in the presence of multiple structural breaks, the approach of Kejriwal and Perron (2008, 2010) was employed in this study. Kejriwal and Perron (2008, 2010) developed an estimation of cointegrated regression models accounting for multiple structural changes. The framework of this approach is general enough to allow for both stationary and nonstationary variables in the model while allowing for serial correlation and heteroskedasticity. The authors illustrated that inference is possible in models with both stationary and nonstationary variables as long as the intercept is allowed to change through regimes. Their work is based on Bai and Perron's (1998) methodology that estimates and tests linear models of stationary variables for multiple structural changes. Kejriwal and Perron (2008, 2010) derived limiting distributions of the sup-Wald test of Bai and Perron (1998) under general conditions for errors and regressors to allow for nonstationary variables in cointegrated regressions.

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

$$y_t = x_t'\beta + z_t'\delta_j + e_t \tag{2}$$

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

$$S_t(T_1, \ldots, T_m) = \sum\nolimits_{i=1}^{m+1} \sum\nolimits_{t=T_{i-1}+1}^{T_i} [y_t - x_t'\beta + z_t'\delta_t]^2 \tag{3}$$

By 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 minima of the sum of squared residuals objective function and can be expressed by the following equation:

$$(\hat{T}_1, \ldots, \hat{T}_m) = \arg\min\nolimits_{T_1, \ldots, T_m} S_t(T_1, \ldots, T_m) \tag{4}$$

The minimization of the sum of squared residuals is obtained in all partitions $(T_1, \ldots, T_m)$, for which $T_i - T_{i-1} \geq q$. The estimates of regression parameters are least-squares estimates associated with partition m $\{\hat{T}_j\}$ — i.e., $\hat{\beta} = \hat{\beta}(\{T_j\})$ and $\hat{\delta} = \hat{\delta}(\{T_j\})$. Bai and Perron (2003) proposed an efficient algorithm for 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. First, the statistics for $UD\max$ and $WD\max$ tests must be calculated. $UD\max$ and $WD\max$ tests are double maximum tests that examine the hypothesis of no structural break against an unknown number of breaks with the given upper bound of breaks M, it and can be calculated by the following formulas:

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

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

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

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

Next, Wald type tests must be applied, where the sup F(0|1) test examines for the hypothesis of no breaks against 1 existing break. If the statistics of this test reject the hypothesis of no breaks, sup $F(l+1|l)$ must be applied to specify the number of breaks in the series. The number of breaks in the series can also be chosen 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 the cointegration tests, the stationarity of the employed variables must be examined. To test the 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. To generate efficient versions of the modified tests of Perron and Ng (1996), Ng and Perron (2001) employed the generalized least squares detrending procedure proposed by Elliot, Rothenberg and Stock (1996). Ng and Perron 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 Akaike information criterion (AIC) and recommended the use of a minimized value of modified AIC for selecting the regression's lag length.

To test the integration properties of variables in the presence of structural shifts, the Carrion-i-Silvestre et al. (2009) test is employed with the null hypothesis of the unit root presence. This test is an extension of the test proposed by Kim and Perron (2009) and allows for up to five breaks at unknown time locations. The Carrion-i-Silvestre et al. (2009) test has the advantage over other alternative tests by allowing structural shifts under both the null and alternative hypotheses. Alternative unit root tests allow structural shifts in the series only under the alternative hypothesis of stationarity, Zivot and Andrews (1992), Perron and Vogelsang (1992), Perron (1997), Vogelsang and Perron (1998).

Before testing the stationarity, the presence of structural shifts in series must be investigated. Ignorance of the presence of structural shifts in a series can lead to misspecification errors. The Perron and Yabu (2009) test investigates for structural changes in the deterministic components of a univariate time series when their integration order is a priori unknown. The F-test has the null hypothesis of no structural shifts and is based on the *Exp* function developed by Andrews and

Ploberger (1994). Three models are estimated by the test, where model I tests for the presence of a structural shift in the level of a variable, model II does so for the slope of the trend, and model III does the same for the level and slope of the trend. Model III, which tests for the presence of a structural shift in both the level and the slope of the time trend, is applied in this study.

2.1. Cointegration

Finally, to test for cointegration characteristics between variables under the consideration of a structural break presence, the Maki (2012) and Carrion-i-Silvestre and Sanso (2006) cointegration tests were employed.

The Maki (2012) test is based on the Bai and Perron (1998) test for structural breaks and the unit root test proposed by Kapetanios (2005). 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 an unspecified number of breaks i that is 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 the cointegration relationships when multiple unknown numbers of breaks exist. When the number of breaks allowed in the Maki test is one, it can be considered as a special case that determines the cointegration test introduced by Gregory and Hansen (1996), which allows for one structural shift. When the number of breaks allowed is two, it presents the special case that coincides with the Hatemi-J (2008) cointegration test in which two structural breaks are allowed.

The Carrion-i-Silvestre and Sanso (2006) cointegration test allows for a structural shift in the cointegrating relationship. The main difference and an advantage of the test over alternative cointegration tests that allow for one structural shift (for example, Gregory and Hansen, 1996) is that it has the null hypothesis of the presence of a cointegration relationship against the alternative hypothesis of no cointegration. Both the null and alternative hypotheses allow for the presence of a structural shift. Allowance of structural shifts in cointegration tests introduces spurious unit root behavior that makes it difficult to reject the hypothesis of no cointegration. Therefore, alternative cointegration tests with null hypothesis of no cointegration have higher a possibility of failing to find the cointegration relationships, which leads to spurious results (Gregory et al., 1996). The Carrion-i-Silvestre and Sanso (2006) test is a Lagrange-multiplier type cointegration test based on the multivariate extension of the Kwatkowski et al. (1992) test. The cointegration test is run for models when the date of the shift is known a priori; when the date is not known, the test estimates the break date by minimizing the sequence of the sum of squared residuals. The estimation of a break date is based on approaches of Bai (1994, 1997) and Kurozumi (2002). Carrion-i-Silvestre and Sanso specified six different models for estimations: model A_n allows for a break in the level, model A has a trend and allows for a break in the level, model B accounts only for a change in the slope of the time trend, and model C allows for a break in both the level and slope of the time trend. Model D allows a break in the deterministic components and the cointegrating vector, and model E contains a trend and allows for a shift in both the deterministic component and the cointegrating vector, similar to model D.

3. Empirical Results

3.1. Unit root tests

To test for the presence of structural breaks in individual variables, the Perron and Yabu (2009) test is employed, see Table 1. The structural break is allowed in both the level and the slope of the time trend of estimated variables.

The null hypothesis of the test, no structural shifts, was rejected for both variables, investment and savings for two estimated periods, 1995–2013 and 2000–2013. The break dates detected by the test are first quarters of 1999 and 2000 for the full period for investment and savings, respectively. These years are characterized by the fast recovery of the Russian economy after the 1998 Russian financial crisis. Between 1999 and 2008, Russia was ranked as one of the world's fastest-growing economies and had the highest per capita income among the BRIC (Brazil, Russia, India and China) countries, which are considered as newly advanced countries (Åslund and Kuchins, 2009). For the 2000–2013 period, the break date according to the Perron and Yabu test was detected as 2008 for both variables, which is characterized by the impact of the global financial crisis. The results of the test demonstrate the presence of structural shifts in estimated time series. Next, the unit root presence in the time series must be estimated. The results of the Perron and Yabu test indicate that the Carrion-i-Silvestre et al. (2009) unit root tests that allow for the presence of structural shifts must be applied to both variables. However, the Kejriwal and Perron (2008, 2010) methodology employed later in this study is designed for cointegrated regression models. Therefore, the standard cointegration test must be applied first, which requires variables to be nonstationary. For this reason, the Ng and Perron (2001) unit root tests are applied first to both variables that do not allow for structural shifts.

Table 2 presents the results of the Ng and Perron (2001) unit root tests. The results are presented for two considered periods, 1995–2013 and 2000–2013. All tests are consistent with one another, and the null hypothesis of the unit root presence was not rejected by any of the tests for either of the employed variables, investments or savings, or for either of considered periods.

Next, the Carrion-i-Silvestre et al. (2009) unit root tests, which allow for up to five structural breaks, were applied to series for both periods. The *t*-statistics of the test and possible break allocations are presented in Table 3. This study allows up to three breaks in the test because results are similar when more breaks are introduced. When structural breaks are allowed, the unit root hypothesis again was not

Table 1
Perron–Yabu test for structural changes in the deterministic components.

Period	EXP-W_{FS} test	$\hat{T}_1$
1995–2013		
Investment	6.38**	2000-Q1
Savings	5.65**	1999-Q1
2000–2013		
Investment	19.50**	2008-Q3
Savings	20.53**	2008-Q3

Notes: ** denotes the rejection of the null hypothesis at the 1% significance level. Trimmer parameter $\varepsilon = 0.15$ is used. The critical values are taken from Perron and Yabu (2009, Table 2c).

Table 2
Unit root tests: Ng and Perron (2001).

Period	Investments				Savings			
	MZ_α^{GLS}	MZ_t^{GLS}	MSB^{GLS}	MP_T^{GLS}	MZ_α^{GLS}	MZ_t^{GLS}	MSB^{GLS}	MP_T^{GLS}
1995–2013								
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 3
Unit root tests: Carrion-i-Silvestre et al. (2009).

Period	MZ_α^{GLS}	MZ_t^{GLS}	MSB^{GLS}	MP_T^{GLS}	$\hat{T}_1$	$\hat{T}_2$	$\hat{T}_3$
1995–2013							
Investment	–29.08	–3.79	0.13	9.83	1996-Q4	2000-Q1	2006-Q3
Savings	–18.08	–2.95	0.16	14.80	1998-Q2	2000-Q2	2008-Q3
2000–2013							
Investment	–15.48	–2.65	0.17	17.97	2002-Q1	2007-Q1	2008-Q3
Savings	–15.74	–2.67	0.17	17.75	2002-Q1	2007-Q1	2008-Q3

Notes: The critical values were obtained by simulations using 1,000 steps to approximate the Wiener process and 10,000 replications. The test is run for model 3, where the structural break affects both the level and the slope of the time trend. Note that for the *MSB* and *MPT* tests, the null hypothesis is rejected in favor of stationarity when the estimated value is smaller than the critical value.

rejected for both periods. The test detected three breaks for every variable for each period. The first break locations are different for the investments and savings series in the 1995–2013 period; for investments, it was the end of 1996, and for savings, it was mid-1998. The end of 1996 for Russia can be characterized by the initiation of negotiations by Russian officials to reschedule the payment of foreign debt inherited from the former Soviet Union. This was the major step toward restoring investor confidence (Chiodo and Owyang, 2002). This shift is characterized by a sharp temporary increase in investments, as seen in Fig. 2. The middle of 1998 for savings is characterized by the effect of the Russian financial crisis that took place in 1998. Meanwhile, 2008 is detected as a break location for both periods, which can be explained by the impact of the global financial crisis.

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 Carrion-i-Silvestre et al. (2009) unit root tests, structural change presence and cointegration tests were conducted.

3.2. Structural change presence

The Kejriwal and Perron (2008, 2010) methodology allows for the presence of nonstationary 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 the variables. Therefore, the Johansen cointegration test was conducted.

To determine the rank of cointegration space, two test statistics are presented, the Trace and the Max-Eigenvalue (Table 4).

The results of the Trace likelihood ratio test statistic and 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 the savings and investment variables for the 1995–2013 period. For the second period, 2000–2013, the estimation results revealed one cointegration equation at the 5% significance level and two cointegration equations at the 10% significance level. Thus, the results of Table 4 indicate the existence of long-run relationships between chosen variables in all cases when structural breaks are not considered.

Having verified the existence of long-run relationships between the variables, the Kejriwal and Perron (2008, 2010) methodology was applied to the series. Table 5 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 5, it can be

Table 4
Standard cointegration test: Johansen.

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

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

Table 5
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)	*UD*max	*WD*max
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 the 5% level. The 5% critical values for the Sup $F(l)$ 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 critical value for the *UD*max test is 14.47. See Kejriwal and Perron (2010). The critical value for the *WD*max test is 9.039. See Bai and Perron (2003). The 5% critical values for the Sup $F(l)$ 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 *UD*max test is 14.79.

Table 6
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 the 5% level; * denotes statistical significance at the 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 $F(l+1 \mid l)$ test are 10.13, 11.14, 11.83 and 12.25 for $l = 1, 2, 3, 4$, respectively, see Bai and Perron (2003, Table 2c).

concluded that there is strong evidence of a structural change present in the employed series in both considered periods.

Table 6 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 the last three columns of the table. In the full period 1995–2013, the sequential test detected two structural shifts, and the BIC and LWZ detected three. In the post-crisis period, 2000–2013, the sequential test did not detect any structural shifts, whereas the BIC and LWZ detected one break. Because the Kejriwal and Perron test (Table 5) provided evidence of a structural shift, the results of BIC and LWZ for one structural shift were considered in this study for the 2000–2013 period.

3.3. Cointegration

Tables 7 and 8 present the estimation results of the cointegration tests that allow for structural shifts. The Bai and Perron (1998) test detected two structural shifts with the sequential procedure and three structural shifts with the BIC and LWZ procedures for the 1995–2013 period, and both procedures detected one structural shift for the 2000–2013 period. First, the Maki (2012) test is applied, which allows for up to five structural shifts in the model; the results for one structural break are consistent with the Hatemi-J (2008) test, and the results for two structural breaks are consistent with the Gregory and Hansen (1996) test.

The results of the Maki (2012) test are shown in Table 7, where MBk presents the t-statistics of the Maki test, and k denotes the maximum number of breaks.

Table 7
The Maki (2012) cointegration test with unknown number of breaks.

Period	*MB*1	*MB*2	*MB*3	*MB*4	*MB*5
1995–2013	–4.70	–7.42*	–8.18*	–9.59*	–9.59*
2000–2013	–4.82	–5.92	–6.51	6.51	–7.05

Notes: * denotes statistical significance at the 5% level. Critical values are taken from Maki (2012, Table 1). The critical values for 1 regressor for trend and regime shifts model with 5% significance level are –5.541, –6.100, –6.524, –7.009, –7.414 for 5 structural breaks respectively for trend and regime shifts model. The trimming parameter is 0.05.

Table 8
Carrion-i-Silvestre and Sanso cointegration test (2006).

Model	1995–2013		2000–2013			
	Test 1	Break 1	Test 1	Break 1	Test 2	Break 2
A_n	0.0741	2006-Q3	0.0526	2006-Q3	0.0526	2006-Q3
A	0.0555	1997-Q1	0.0455	2006-Q3	0.0455	2006-Q3
B	0.0506	1997-Q3	0.0871	2009-Q1	0.0822	2006-Q3
C	0.0614	2000-Q1	0.0407	2006-Q4	0.0459	2006-Q3
D	0.0750	2006-Q3	0.0482	2006-Q3	0.0482	2006-Q3
E	0.0570	2000-Q1	0.0370	2006-Q3	0.0370	2006-Q3

Notes: Test 1 is the test when the break date is a priory unknown, and the date location is determined by the test. Test 2 provides statistics when the break date is known and exogenously is determined.

The test statistics rejected the null hypothesis of no cointegration for the 1995–2013 period when more than one break is allowed. However, when one break is allowed, it failed to reject the null. The test statistics did not detect cointegration relationships for the 2000–2013 period for any number of structural shifts allowed. Based on the results of the Bai and Perron (1998) test, the 1995–2013 test is characterized by the presence of two or three structural shifts. In the presence of these shifts, the Maki test provided evidence of the cointegration relationships for the 1995–2013 period. The 2000–2013 period is characterized by the presence of only one structural shift; however, the Maki test statistics failed to provide evidence of cointegration for this period.

Therefore, the Carrion-i-Silvestre and Sanso (2006) cointegration test is applied, which is more powerful compared with alternative tests in finding evidence for cointegration relationships in the presence of a structural shift. The test is applied to the model for both periods for reasons of comparison. The null hypothesis of cointegration could not be rejected by any of 6 considered models at the 5% significance level for both estimated periods. Therefore, it can be concluded that long-run relationships between investment and savings exist in both estimated periods when they are affected by the presence of structural shifts. The Carrion-i-Silvestre and Sanso test investigates the estimated model for the presence of cointegration when a priori unknown and known break dates are allowed. Test 1 in Table 8 presents the test statistics when the break location is a priori unknown, and Break 1 is the date for the structural shift, which is estimated by the test. Test 2 shows estimation results when the break location is exogenously determined. The period 1995–2013 is characterized by the presence of three structural shifts (Table 6); therefore, only Test 1 was applied to this period for the comparison. The period 2000–2013 is specified by one structural shift (Table 6); therefore, test 1 and test 2 were applied. Test 2 is applied to the model in which the break date is determined at the 2006-Q3 location, which is detected by the Kejriwal and Perron test (Table 9). In the 1995–2013 period, the break date is estimated by testing at different locations for different models, which indicates that the period has more than one structural shift. The test estimations for the 2000–2013 period detected break 1 in five out of six models at

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	20.13**	15.85**	18.36**	21.28**	1997-Q3	2000-Q2	2006-Q3
	(0.03)	(0.87)	(0.93)	(1.01)	(0.95)	('96-Q4–'98-Q1)	('99-Q3–'02-Q2)	('06-Q1–'07-Q1)
(S)	0.05	18.32**	15.47**	19.24**	–	1997-Q3	2006-Q3	–
	(0.04)	(1.08)	(1.21)	(1.17)		('95-Q3–'97-Q4)	('06-Q2–'11-Q3)	
2000–2013								
(BIC, LWZ)	–0.10**	21.17**	24.02**			2006-Q3		
	(0.03)	(1.12)	(1.03)			('05-Q3–'07-Q1)		

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

the same location, which is determined by the Kejriwal and Perron test (Table 9), it is 2006-Q3. The Kejriwal and Perron test detected this particular break date for the 1995–2013 period as well (Table 9). The shift in investment—savings relations can be explained by fast-growing private sector debts based on government policies to stimulate foreign capital inflows that can be invested better than domestic capital (Gaddy and Ickes, 2010). The policy that stands behind these changes is the lift by the Russian government and the Bank of Russia of almost all restrictions on capital transactions. Explosion in foreign capital inflows led to substantial growth in lending (Kudrin and Gurvich, 2015).

The results of the cointegration estimations that allow for structural shifts provide strong evidence for the existence of cointegration relationships in both periods. In the literature, the cointegration 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 existence of long-run relationships in the presence of structural breaks supports the solvency existence of a current account in Russia in both considered periods.

3.4. *Coefficients estimates*

Table 9 reports the results of the parameter 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 savings retention coefficient, $\hat{\beta}$, corrected for the presence of structural breaks, are given in the second column. Break locations detected by the Kejriwal and Perron test are consistent with break locations detected by the Carrion-i-Silvestre and Sanso test (Table 8). The first break is detected at 1997-Q3, which is characterized by the external shock of the Asian financial crisis. The second shift is estimated at 2000-Q2, which is characterized by the fast recovery after the Russian financial crisis of 1998. All three procedures—S, BIC and LWZ—determined a common 2006-Q3 break location for both estimated periods 1995–2013 and 2000–2013, which can be explained by the increase in private debt and foreign capital inflow.[4]

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 savings retention coefficient estimates were not found to be significant. Estimations of the post-crisis period 2000–2013 produced significant results for the savings retention coefficient when one structural break was detected by the BIC and LWZ procedures. Thus, the estimate of the savings retention coefficient in the presence of a structural break was found at the level –0.10, which is relatively close to zero.

[4] A detailed analysis of this period is in the explanation part of Table 8.

Table 10
Estimated regression parameters OLS and FMOLS.

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

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

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 the coefficient estimations with allowance for structural shifts. The savings retention coefficient was found to be insignificant in the full considered period, 1995–2013. However, the estimations for the post-crisis period revealed a significant savings retention coefficient with a negative sign at the –0.275 level and the –0.306 level by the OLS and FMOLS procedures, respectively. The negative sign of the savings retention coefficient can be interpreted as a low correlation between saving and investment flows or as the existence of high savings flight abroad owing to a deficiency in the domestic financial structure (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 a foreign currency, and to transfer financial assets in a national currency into financial assets denominated in a foreign currency. Domestic capital flight has existed since the Russian economy moved to the market economy model. However, capital flight from Russia is mainly not connected to the normal decision of profit maximization; 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 approximately one-quarter of inward FDI was a result 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 composing 4% of the GDP can be explained by an unfavorable investment climate. Capital flight from Russia, $133.7 billion in 2008, decreased to $56.1 billion in 2009 and to $34.4 billion in 2010 and then rose to $80.5 billion in 2011 and $56.8 billion in 2012.[5] The main concern of domestic capital outflow in Russia is its effect on domestic investments; therefore, 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 on this list in 2012.[6] As a result, the level of capital mobility has continuously increased in Russia, reducing the level of correlation between investments and domestic savings.

[5] Sergei Ignatyev, Chief of the Central Bank of Russia. RIA Novosti, 2013, June 5.
[6] UNCTAD, Global Investment Trends Monitor.

The results of the saving retention coefficient estimates illustrate a high mobility of capital in Russia in the post-crisis period. Consideration of structural shifts does not significantly affect estimation results in which 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 for the presence of the FHP 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 by employing panel econometric techniques such as the Fully Modified OLS (FMOLS), Dynamic OLS (DOLS), and Pooled Mean Group (PMG). However, each panel cointegration estimation method provided different results for the individual countries. Thus, the savings retention coefficient for Russia was found to be significant in all three cases, but the values were found at different levels (–0.21, –0.02, and 1.49, respectively) depending on the method employed. Therefore, it is difficult to draw a certain conclusion without choosing a particular method. Trunin and Zubarev (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 savings retention coefficient for the period 1996–2011 in Russia was not found to be significant at the 0.221 level, which is compatible with the present study results. In the post-crisis period 2007–2011, the savings retention coefficient was found to be significant at the 0.8 level, indicating a low capital mobility level after the global crisis. However, the latest estimations considered only five observations, which is not sufficient 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 for the presence of the FHP in Russia in the post-crisis period, whereas estimations with accommodation for structural breaks illustrate high capital mobility and no evidence of the FHP.

4. Conclusion

This paper examined capital mobility in Russia in the presence of structural breaks for two periods: 1995–2013 and the post-crisis period from 2000–2013. Recently developed econometric methods were applied to quarterly series to investigate the cointegrating relationships of investment and savings variables, considering the presence of structural shifts in the model when relevant, and to estimate the savings 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, 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 present in the 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 were detected by the BIC and LWZ procedures. In the post-crisis period, 2000–2013, one shift was detected by both the BIC and LWZ procedures.

To examine the cointegration relationships of the series in the presence of structural breaks, the Maki (2012) and Carrion-i-Silvestre and Sanso (2006) cointegration tests were employed. The Maki test allows for the presence of possible multiple breaks and has a null hypothesis of no cointegration. The Carrion-i-Silvestre and Sanso test allows for the presence of one structural shift and has a null hypothesis of cointegration. The results of the Maki test provide evidence of the existence of cointegration relationships in the 1995–2013 period when more than one break is allowed. The Maki test did not provide evidence of cointegration for the post-crisis period. Therefore, the Carrion-i-Silvestre and Sanso test was applied, which did not reject the null hypothesis of cointegration for any estimated period, providing strong evidence of cointegration relationships in the model when affected by a structural shift. Existence of long-run relationships with the introduction of structural breaks indicates the solvency of a current account in Russia in both of the considered periods.

The OLS and FMOLS estimates of the savings 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 to be significant in the full estimated period, 1995–2013. However, estimations of the post-crisis period were found to be significant with a negative sign at the –0.275 and –0.306 levels by the OLS and FMOLS procedures, respectively, and at the –0.10 level 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 savings 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 illustration of better capital mobility. Thus, the results of this study employing OLS and FMOLS estimations provide weak evidence of the FHP in Russia in the post-crisis period, whereas estimations with accommodation of structural breaks illustrate high capital mobility and no evidence of the FHP.

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