КОНСОРЦИУМ ЭКОНОМИЧЕСКИХ ИССЛЕДОВАНИЙ И ОБРАЗОВАНИЯ
ECONOMICS EDUCATION AND RESEARCH CONSORTIUM
Working Paper Series

ISSN 1561-2422

COVERED INTEREST PARITY: EVIDENCE FROM RUSSIAN MONEY MARKET

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Working paper No 16/01E

This project (No. 13-5771) was supported
by the Economics Education and Research Consortium
and funded by GDN

Project No. 13-5771

Research Area C

Covered interest parity: evidence from Russian money market»

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JEL Codes: E43, F31, G15.

Abstract. This paper tests covered interest parity at Russian money market over period of 2010-2014 and studies scale and sources of deviations from it. We use both offered and actual interbank interest rates for four different terms. Average deviations from the parity vary between 8 and 105 basis points depending on rates and terms. We test credit risk, turbulence and monetary policy as explanation of these deviations and assessed them quantitatively. For example, one standard deviation change in credit risk is responsible for 50 per cent of the average deviation from parity compared to 72 per cent due to monetary policy spread and (minus) 22 per cent due to turbulence for one week offered rate spread. Risk and turbulence effects grow with maturity and higher for actual rate spreads.

Acknowledgements. We are grateful to EERC experts and discussants, especially to Tigran Poghosyan, Natalia Volchkova, and Diana Weinhold for their insightful comments and useful suggestions enriched significantly our research. We also thank Vladimir Pyrlik for his helpful comments. The gratefully acknowledge financial support of EERC under the grant No. 13-5771.

Key words: Covered interest rate parity, CIP, Russian money market, credit risk, turbulence.

Contents

1. Introduction

Interest rates reflect the tightness of monetary policy. Central banks use them as intermediate targets to achieve certain macroeconomic conditions. According to covered interest parity (CIP) theory, a country's domestic interest rate is linked closely with external rates. To conduct monetary policy, central banks have to take this parity into consideration. Slight deviations from CIP, indicating a looser link between the domestic and the foreign capital markets, are nonetheless typically observed. Saito and Shiratsuka (2001) argue that deviations from parity indicate liquidity constraints in the banking sector, which leads to limited arbitraging and money market segmentation. As a result, lending activity decreases. A breakdown in transmission mechanisms of this kind has been noted recently by a number of researchers and practitioners. Kovalenko (2010), among others, stresses high segmentation in the Russian interbank market.

The purpose of our project is to estimate the presence, scale and causes of deviations from CIP in the Russian money market.

Ascertaining whether systematic deviations from parity take place is not as easy as it may seem. One of the issues here is choosing proper interest rate measures. It is typical to use offered rates as indicators of money market conditions. During the recent financial crisis, however, it turned out that offered rates might be confusing, even when related to the most developed markets. Dollar LIBOR is a good example (Michaud and Upper, 2008). Until recently, little attention was paid to actual money market rates, as the offered rates behaved well. The Lehman Brothers collapse and subsequent market turmoil showed, however, that offered rates represent market rates poorly, at least in periods of turbulence. The actual rates are determined bilaterally and there is no statistics on these rates for the world's most important money markets. In contrast, the Central Bank of Russia compiles actual as well as offered rates and publishes them regularly, though in aggregate form. The estimations of and explanations for deviations from parity based on these rates may produce either similar or qualitatively different results. The latter case is a sign that offered rates may be unreliable. The obtained results thus provide additional insight into discussions concerning money market indicator choices.

While deviations from parity may attract the attention of monetary authorities, it is unclear why they happen, therefore it is also difficult to understand how to react to them. Intending to clarify the nature of the deviations from parity, we examine three possible sources of such deviations, namely transaction costs, credit risk, and monetary policy.

2. Literature review

2.1. Covered interest parity (CIP)

CIP hypothesis is not new in economics. The development of the forward foreign exchange (FX) market at the end of the 19th century put covered interest rate arbitrage into textbooks for practitioners (Dent, 1920). Keynes (1924) discussed both parity condition and sources of deviation from it. A basic covered parity relation means equality of yields on comparable foreign and domestic assets, which could be written as:

$$1 + i_t = \left(1 + i*_y\right)\frac{F_t}{S_t}, \qquad (1)$$

where i_t is an interest rate at domestic currency market, $i*_t$ is a foreign currency market interest rate, S_t is a spot foreign exchange rate and F_t is a forward foreign exchange rate. It is assumed that forward and credit maturities are the same and interest rates are not annualized.

Interest rate differential (d_t) and forward premium (f_t) are described by (2) and (3). In words, CIP thus means the equality of the interest rate differential to the forward premium.

$$d_t = \frac{1 + i_t}{1 + i*_y} \approx i_t - i*_y, \qquad (2)$$

$$f_t = \frac{F_t}{S_t} - 1 \approx \ln(F_t) - \ln(S_t). \qquad (3)$$

The parity occurs because of arbitrage. Let us assume that the foreign interest rate is lower than the domestic one: $i > i^* + f$. In a risk-free economy without transaction costs, this deviation creates arbitrage opportunities, as shown in Fig. 1 below.

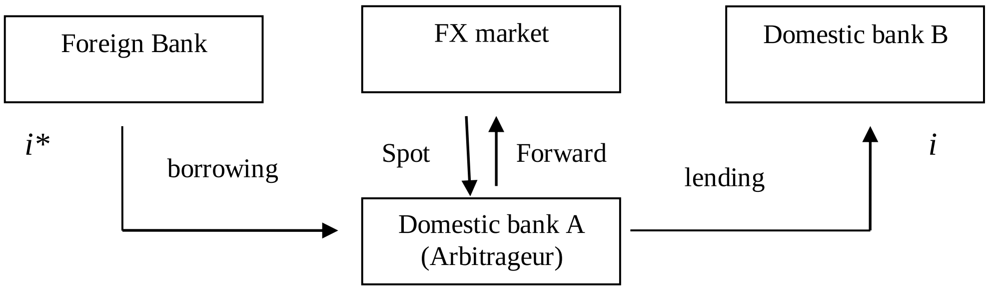

Fig. 1. Covered interest arbitrage mechanism

By borrowing at *i** from the Foreign Bank and lending to Domestic Bank B at *I*, the Arbitrageur obtains profit $\pi = i - i^* - f$ per one domestic currency unit.

Several approaches to testing CIP empirically have been applied by researchers. The first method is based on checking if actual deviation from parity (the interest rate differential minus forward premium), differs from zero (Taylor and Tchernykh-Branson, 2004; Takezawa, 1995; Taylor, 1989; Fletcher and Taylor, 1994; Batten and Szilagyi, 2010).

The second approach for testing CIP is based on regression of the interest differential on the forward premium (4) or, alternatively, regression of returns on assets in domestic currency on the returns in foreign currency corrected for the forward premium (5). The equations (4) and (5) are virtually the same if the foreign interest rate (i^*_t) is lower, which is typically the case. The choice of model depends mainly on data availability.

$$d_t = \alpha + \beta \frac{F_t}{S_t} + \varepsilon_t, \qquad (4)$$

$$1 + i_t = \alpha + \beta\left(1 + i^*_y\right)\frac{F_t}{S_t} + \varepsilon_t \qquad (5)$$

If CIP holds, α and β should be equal to zero and unity respectively in both cases. Equation (4) has been analyzed by many authors for various data and time horizons. Older studies (Branson, 1969; Popper, 1993) did not pay attention to the time-series nature of the data. Maenning and Tease (1987) allowed for autocorrelation, but did not test for co-integration. Moosa and Bhatti (1996) and Gurvich et al. (2009) found cointegration between the returns on domestic and foreign assets. The restricted model was, however, rejected in the first paper, while in the latter one the test was not reported.

Al-Loughani and Moosa (2000) suggested a third approach to testing for CIP. They tested variances of differently hedged portfolios for equality instead of testing means. Authors reported on holding CIP and agreement between the latter data and cointegration-based approaches.

Papers concerning CIP in the Russian money market include Gurvich et al. (2009); Kovalenko (2010); Taylor and Tchernykh-Branson (2004); Skinner and Mason (2011).

To investigate the effect of the the Central Bank of Russia's exchange rate policy on deviations from CIP, Gurvich et al. (2009) conducted cointegration analysis. They used daily non-delivered forward (NDF), LIBOR and Moscow Interbank Offered Rate (MosIBOR) data for 2001 to 2008. Published estimates show CIP violation for at least two of three sub-periods. Gurvich et al. (2009) argue, however, in favor of CIP, as cointegration between the rates is observed. Their extremely weak CIP hypothesis seems to be explained by focusing mainly on the consequences of FX market liberalization.

Kovalenko (2010) regressed the forward premium on interest rate differential i-$i*$ between the Russian and London money market rates. She concludes that the relationship between rates was close to CIP, based on monthly averages of LIBOR and domestic offered actual and bid rates (MIBOR, MIACR, and MIBID respectively[1]) with one month maturities for the period between 2001 to 2010. Unfortunately, corresponding test statistics are not published.

[1]MIBOR is Moscow interbank offered rate, MIACR is Moscow interbank actual credit rate, MIBID is Moscow interbank bid rate. All these rates are calculated by the Central Bank of Russia.

Taylor and Tschernyh-Branson (2004) estimated the threshold autoregression (TAR) model for Russian and U.S. treasury bills for the period from December 1996 to August 1998. They reported violations of CIP and that the asymmetric neutral band shifted upward. Their analysis concerned a period of high instability in Russian financial markets. The authors therefore attributed their findings to risk premiums.

Close results were obtained by Skinner and Mason (2011). They used daily data for the period from January 2003 to October 2006 from a panel of countries. They didn't reject the CIP hypothesis for Russia concerning five-years and three-month maturities. In fact, they found that the average deviation from CIP was less than one basis point for three-month maturities. They did, however, obtain non-stationary GARCH residuals for the latter maturity.

As we can see, several papers address CIP in Russian money and debt markets, but they don't provide exhaustive discussion of this issue.

The next important question concerns the persistence of the profitable opportunities. While earlier papers such as Fletcher and Taylor (1996) and Taylor and Tchernykh-Branson (2004) reported that long-lived profitable opportunities exist, recent studies demonstrate clearly the opposite. Akram et al. (2008), for example, investigated the properties of potential departures from CIP conditions. They conclude that duration of CIP deviations didn't seem to last more than a few minutes. In most cases, the average duration was between 20 seconds and 4 minutes.

2.2. Explaining deviations from CIP

All of the previous investigations into covered interest parity in the Russian money market, with the exception of Skinner and Mason's (2011), do not focus on the question of why disparities exist and how to explain them. We intend to fill this gap with our own research.

Arbitrage from foreign to domestic currency involves both transaction costs (i.e. the cost of screening the borrower for reliability), *t*, and the price for the default risk of domestic bank B (see Fig.1) (Akram et al, 2008). In this case, CIP arbitrage is not profitable under condition (6):

$$i < i^* + f + t + r \qquad (6)$$

Noticeably, we have different kinds of transaction costs and risk premiums depending on arbitrage direction. Arbitrage from domestic to foreign currency is accompanied by costs to screen foreign borrowers for reliability (t^*) and a price for default risk (r^*). Transaction costs and risk premiums vary from one transaction to another due to the heterogeneity of borrowers. No-arbitrage condition for the case of $i_t < i*_t + f_t$ is therefore presented by equation (7):

$$i + t^* + r^* > i^* + f \qquad (7)$$

Inequalities (6) and (7) imply the absence of arbitrage opportunities when deviations from parity are small compared to transaction costs. Frenkel and Levich (1977) derived the costs from either bid-ask spread or brokerage fees. They conclude that transaction costs captured at least 85% of the apparent profit opportunities originating from the deviation from CIP.

Clinton (1988) estimated the no-arbitrage band due to transaction costs for the U.S. dollar and five most traded currencies as ±0.06% per annum and showed that profitable arbitrage opportunities are not infrequent. Fletcher and Taylor (1996) came to similar conclusions based on data concerning long-term contracts. Transaction costs prevent arbitrage when deviations from parity are small. It is therefore natural to expect more persistent deviations from CIP within the neutral band than outside of it. Taylor and Tchernykh-Branson (2004) and Hutchison et al. (2012) used this difference to estimate the no-arbitrage band.

The recent financial crisis attracted attention to counterparty risks affecting CIP. To capture these risks, Levich (2012) used CDS; Baba and Packer (2009a; 2009b) exploited the LIBOR-OIS spread, while Skinner and Mason (2011) chose VIX volatility index and changes in the slope of the Treasury yield curve as proxies for this risk. All these studies found that this risk significantly influenced deviations from CIP. Hui et al. (2011) also stressed the strong effect of the LIBOR-OIS spread on deviations from CIP during the recent financial crisis. Taylor and Tchernykh-Branson (2004) studied spreads between Russian and U.S. short-term bonds just before August 1998. They attributed the asymmetry of the no-arbitrage band to the spreads of risk premiums. This explanation seems reasonable taking into account the fact that a foreign debt moratorium was announced by the Russian authority in August 1998. Taylor and Tchernykh-Branson (2004) noticed a possible avenue for future research, including extension of the length of time over which data is collected and explaining Russian default premiums by making them dependent on other variables.

There is some evidence that deviations from CIP depend on turbulence in financial markets. Branson (1969) and Taylor (1989) noted that deviations from parity grow with turbulence. Skinner and Mason (2011) also attributed deviations from CIP in the Russian money market to instability in financial markets. They used several indicators, including TED spread, CDS spread, VIX, and equity premium to capture the effect of financial turmoil on deviations from CIP.

Aliber (1973) stressed political risk as a reason for CIP disparity. Assets available in money markets differ in two ways, firstly, in the currency in which they are denominated and secondly, in the political jurisdiction in which they are issued. These differences lead to exchange and political risks, the first of which can be eliminated by purchasing forward contracts. Consequently, the existence of political risk can potentially influence CIP conditions. Empirical analysis has indicated that deviations from CIP are smaller among deposits in different currencies within one jurisdiction than among deposits in different currencies in different jurisdictions. Using Aliber's concept, Dooley and Isard (1980) estimated a model of portfolio behavior and found out that political risk influences the interest rate differential. They collected Euromark rates and interest rates on mark-denominated loans in Germany during the early 1970s, when Germany imposed controls on the inflow of capital several times. The results of this estimation showed that most of the fluctuations of the interest were caused by capital control.

Monetary policy affects money market rates, but it has no place in the basic parity relationship. As a result, changes in money rates should be fully offset by currency depreciation rate shifts. The Central Bank of Russia uses a corridor-type monetary policy, as does the

European Central Bank. Several papers, including Wurtz (2003) and Beirne (2012), discuss the negative effect of providing additional liquidity on overnight money rates. Egorov and Kovalenko (2013) demonstrate the same effect on the Russian money market. Abbassi and Nautz (2012) show that monetary surprises have a significant effect on Euro Interbank Offered Rate (EURIBOR).

We conclude that the evidence on presence and sources of deviations is ambiguous. CIP studies related to Russia do not provide exhaustive discussion of this issue. Moreover, all of the studies except Skinner and Mason's (2011) are not focused on the question of why such disparities exist. Latter papers have attempted to explain the deviations using VIX, equity premium, and the yield curve slope, but these factors have turned out to be insignificant, prompting us to further examine the issue.

3. The Russian money market and monetary policy framework

3.1. Russian money market survey

Dominant large banks and a dispersed competitive fringe are striking features of the Russian banking system. While the biggest banks often have their roots in Soviet special banks (Sberbank, VTB, VTB24) or their reincarnations (Rosselkhozbank), numerous smaller banks were founded, mainly in the late 1980s and early 1990s when banking became both a key business service and a suitable device for rent-seeking. The total number of banks has been decreasing for about two decades. It dropped from 1078 to 992 between 1 April 2010 and 1 1 2014. The five largest banks, however, control about half of the total assets within the banking system and the assets of the 200 largest banks comprise approximately 95% of the total assets in banking system (see Table 1).

Table 1. Russian Banking system. Selected indicators.

Credit institutions by total assets	Total assets			
	1.03.2011		1.03.2014	
	bn rubles	%	bn rubles	%
Top Five	16,381	48.4	31,769	53.7
6-20	6,814	20.1	10,983	18.6
21-50	3,969	11.7	6,573	11.1
51-200	4,661	13.8	6,988	11.8
201+	2,032	6.0	2,825	4.8
Total	33,858	100.0	59,137	100.0

Note: Due to rounding, some totals may not correspond with the sum of the separate figures.
Sources: *Bulletin of Banking Statistics* (various issues); Table 4.1.5; authors' calculations.

It is commonplace in Russian literature that the ownership of banks significantly affects their market behavior. State-controlled banks accumulate more than half of the assets within the banking system (Vernikov, 2009; Kovalenko and Kislyakova, 2010). The remaining share of the assets is divided between the national banks and foreign capital-controlled banks in approximately equal shares.

Kovalenko and Kislyakova (2010) provided a survey of banks' behavior in the interbank market between 2007 and 2009 based on monthly balance sheets. According to them, 90% of interbank money market turnover, 85% of interbank borrowings and 87% of interbank borrowings from abroad were made by the 73 largest Russian banks (as of December 2009). These numbers are, however, comparable with their share of the banking system’s total assets. About 300 banks were inactive, with lending and borrowing in the interbank money market accounting for less than 1% of their assets.

Before the recent global financial crisis, Russian banks borrowed extensively from the international interbank money market. According to Kovalenko and Kislyakova (2010), net external borrowings peaked in August 2008 at a level of US$ 76 billion. Later, the state-controlled banks as well as the private national banks started to substitute these foreign borrowings with domestic sources. In particular, state-controlled banks and other biggest and most reliable banks received loans from the central bank.

Figure 2 presents the percentage of banks' assets that were comprised of central bank loans as of 1 July 2014. As we can see, the largest banks used this resource much more intensively, while the smallest ones relied on it least.

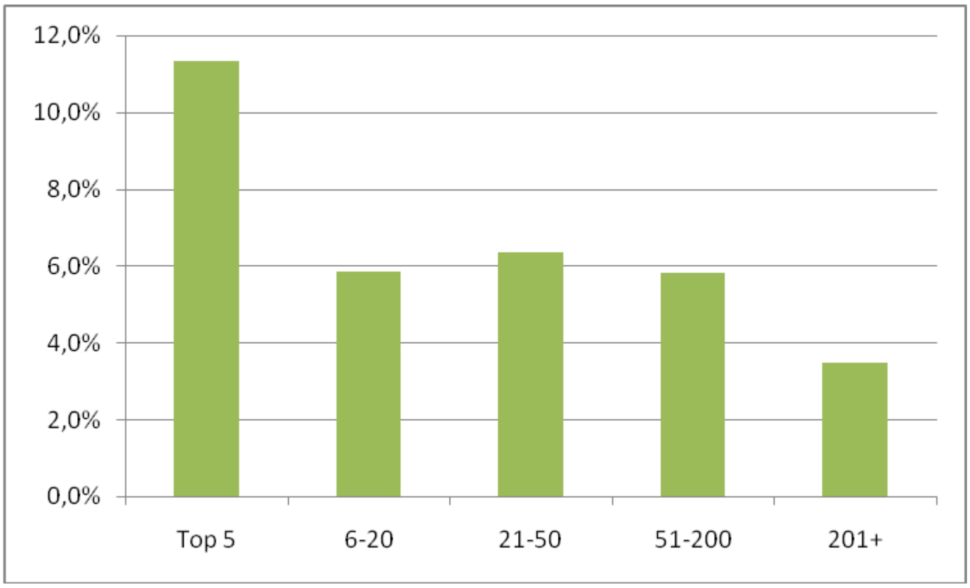

Fig. 2. Loans from Central Bank of Russia in assets of banks by bank size.
Source: *Bulletin of Banking Statistics*, 2014, no. 7, table 4.1.1.

Turning to domestic capital reshaped the internal interbank money market. It became more liquid; larger banks started to operate more actively, lending to some riskier banks and absorbing the abundant liquidity of others. The terms of internal interbank borrowings went up. Another striking change was a rapid growth in borrowing from the central bank.

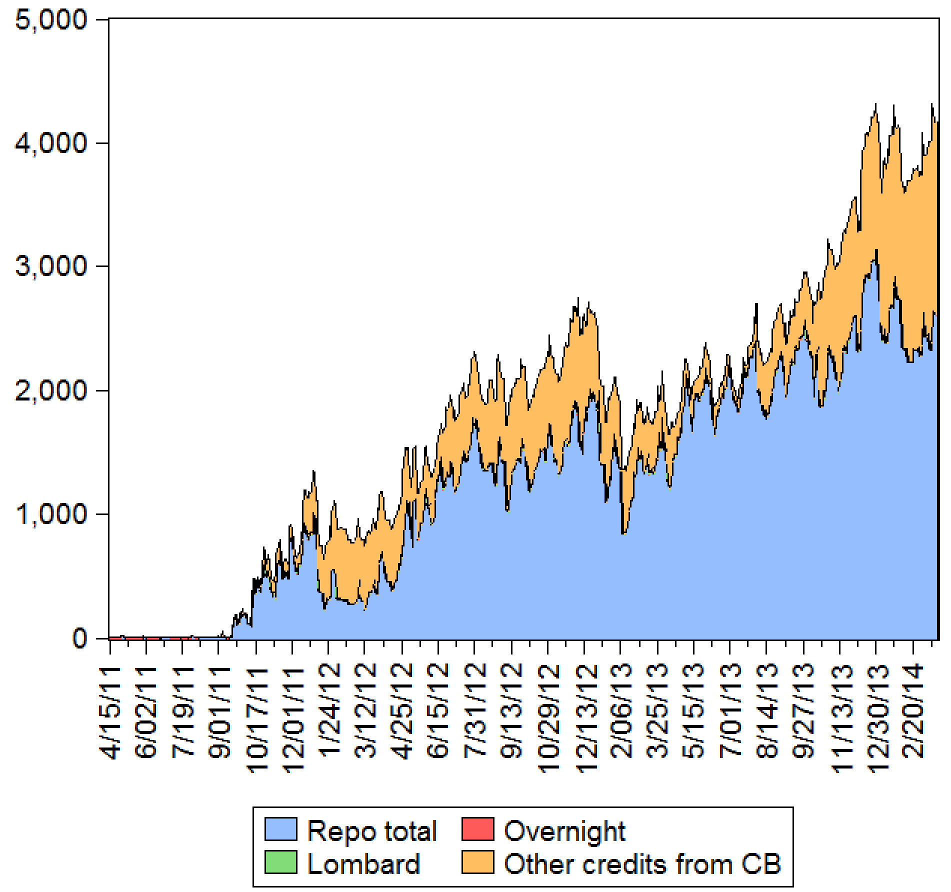

Fig. 3. Loans from the Central Bank of Russia in assets of banks by type of loan
Source: Central Bank of Russia

Figure 3 presents the dynamic and structure of banks' borrowings from the central bank.

Whilst in the first half of 2011 the borrowings were insignificant, they had reached 7% of the total assets of banks by early 2014.

REPO operations were the main type of borrowing from the central bank. Other credits (or in other words loans) backed with less liquid assets were becoming more important as total borrowings grew and banks lacked liquid collateral. Overnight and Lombard credits were insignificant.

3.2. The changing Russian monetary policy framework

The Russian monetary policy framework has shifted gradually during the last decade. For many years, the Central Bank of Russia has paid attention both to smoothing exchange rate fluctuations and stabilizing inflation. The relative importance of these goals, however, has varied over time.

The central bank strictly targeted the price of a bi-currency basket composed of $0.55 and €0.45 before the global financial crisis. The sudden shift in the terms of trade in late 2008 led to a rapid devaluation and to an enormous widening of the pegged corridor up to 15 rubles, or approximately ±20%.

As the acute phase of the crisis passed away, the central bank introduced a new corridor and a corridor adjustment rule. The corridor was three rubles wide and the prescribed currency intervention scale was dependent on the deviation of the current basket price from its middle point. This middle point was changing linearly with cumulative interventions. The rule had two parameters; the corridor width describing the foreign exchange rate volatility in the short term and the sensitivity of the middle point to interventions related to the speed of the exchange rate adjustment over the long term.

During the studied period, the central bank gradually relaxed the foreign exchange rate peg. The corridor width had been set at 5 rubles since 1 March 2011, for example, while every $600 million in cumulative interventions shifted the midpoint for 0.05 rubles. The corridor width was later extended up to 7 rubles, while the cumulative interventions necessary to adjust the midpoint for 0.05 rubles were lowered to $350 million.

While exchange rate policy was changing to become freer floating, the interest rates were pegged increasingly tighter. During this period, a corridor-type monetary policy framework was developed. Deposit standing facility rate forms the lower bound of the corridor. The upper bound of the corridor is determined by the REPO standing facility rate. In April 2011, the deposit standing facility rate was 3%, while the REPO rate stood at 6.75%. By the end of 2012, the corridor width had shrunk gradually to 2 p.p. and was fixed at that level.

During the periods of structural liquidity deficit, the central bank injected a lion's share of liquidity through floating interest rate auctions. This is a 'pay-as-bid' auction, with the reserve interest rate typically set in the middle of the corridor. This rate has served as a key policy rate since September 2013. The central bank determines the liquidity amount to offer at the auction based on its banking sector liquidity forecast and lends money through floating-rate REPO auctions for 1-6 days, for one week and occasionally for longer periods. Emphasis was shifted from daily to weekly REPO operations. Weekly auctions became a main refinance operation for the central bank up to the end of 2013, while shorter loans retained the function of fine-tuning during periods of short-term liquidity imbalances.

Lack of marketable collateral may prevent banks from borrowing through the REPO mechanism. To provide these banks with liquidity, the central bank uses longer-term credits secured with illiquid assets. These credits have been used most actively during periods of severe deficit of liquidity in the market, particularly at the end of the studied period.

To summarize, the Russian monetary policy framework is very complex and has changed gradually during the studied period. The central bank shifted its focus from the foreign exchange rate to the money market rate as an intermediate target. A set of interest rate instruments for providing and absorbing liquidity created a corridor-type framework. The central bank uses not only price instruments like standing facilities, but also quantity instruments such as floating-rate REPO auctions.

3.3. Monetary policy and deviations from CIP: A model

To trace the effect of monetary policy on money market rates, we propose a simple model. Let commercial banks be distributed by marginal benefits from their liquid assets i and risk premiums associated with their borrowings rp. Each bank is endowed with liquidity asset stock h. These assets could be placed in the central bank as deposits at rate d or lent to another bank j at a rate $l_j = i^* + rp_j$, where l_j, is a risky interbank rate for bank j, i^* is a risk-free money market rate and rp_j is a risk premium associated with the borrowing bank. Aside from borrowing from the interbank market, banks may switch to the central bank's secured loans (REPO) at rate r_x which is the same for all banks. For simplicity in both cases, the borrowings are limited to a.

In this model we assume that banks are distributed by i and rp uniformly with density $f(rp,i)=(\overline{rp},\bar{i})^{-1}$ if $rp\in[0,\overline{rp}]$ and $i\in[0,\bar{i}]$.

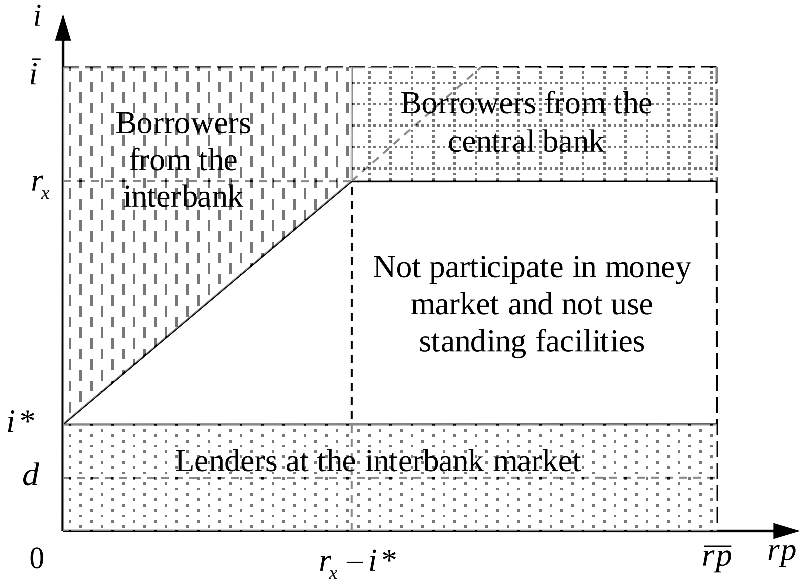

Fig. 4. Banks' participation in the money market

Fig. 4 reflects banks' participation in the interbank market and their usage of the central bank's standing facilities depending on i and rp.

We derive interbank market liquidity demand and supply functions and solve the model for risky and risk-free rates and transaction volume as well as for demand for lending from the central bank and supply of deposits in the central bank. These results are presented graphically in Fig. 5, while derivations and analytical results are provided in Appendix 2.

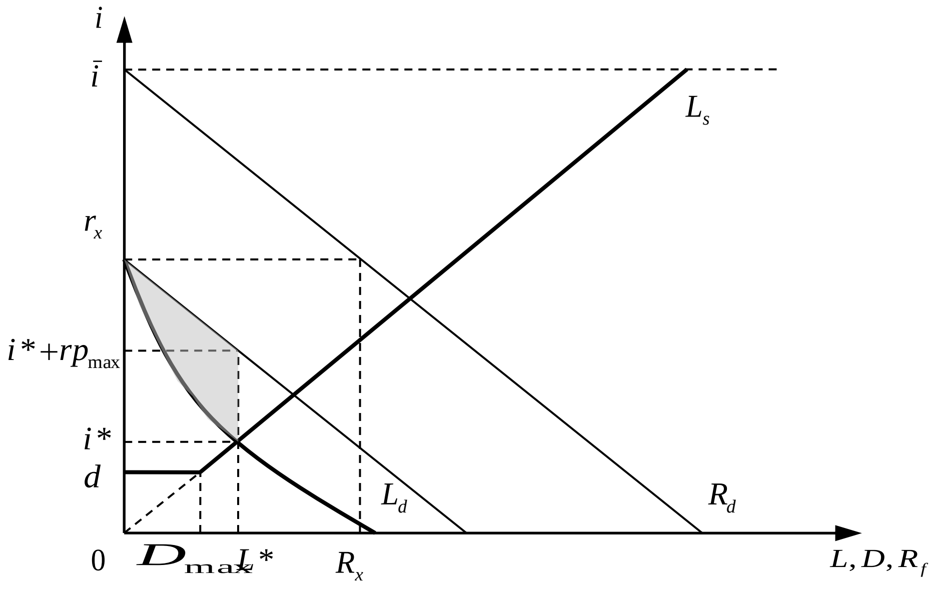

Fig. 5. Money market equilibrium

The thicker lines L_d and L_s reflect liquidity demand and supply schedules correspondingly. The thin line R_d is the demand for secured central bank lending. Given the central bank's rates r_x and d equilibrium transactions volume L^* the risk-free rate i^* and borrowings from the central bank R_x are shown in the diagram as well. The total risk premium in the money market is shadowed.

Obviously, higher lending rates increase both risky and risk-free rates. The risky rate, however reacts more strongly, as riskier banks turn from costlier secured borrowings from the central bank to cheaper finance from the banking sector (see Appendix 2). As a result, a higher interest rate on central bank funding can be associated with higher deviation from parity due to risk premiums. We use the lowest interest rate on actual borrowings at floating-rate auction as a measure of the central bank lending rate r_x.

4. Data

We have used freely available data published by the Central Bank of Russia and National Foreign Exchange Association (NFEA). Table 2 presents the time-series. Descriptive statistics are in Appendix 3.

Table 2. Russian money market indicators

	MIACR-IG	Mosprime	Implied rate	ROISfix	RTSVX
Maturities[2]	from 2 to7d, from 8 to 30d, from 31 to 90d, from 91 to 180d	1w, 1m, 3m, 6m	1w, 1m, 3m, 6m	1w, 1m, 3m, 6m	N/A
Contributor banks	Counterparty banks of transaction	minimum 8 contributors	8-18 contributors	minimum 6 contributors	N/A
Publisher	CBR	NFEA			Moscow Exchange
Calculation methodology	a weighted average of actual interest rates on interbank loans	a simple average using all rates given by contributors except for the two lowest and two highest rates		a simple average of midpoints between bids and asks given by contributors except for the two lowest and two highest rates	Methodologyp rovided by Moscow Exchange[3]
Missings	Yes	No	No	No	No
Period	01/04/2010 – 03/03/2014			15/04/2011 – 03/03/2014	24/01/2011 – 03/03/2014

[2] d, w, and m denote day, week, and month correspondingly.
[3] See http://fs.moex.com/files/5091.

All the data is published on a daily basis. Our sample includes daily observations since 1 April 2010, as there is no earlier data on the ruble-denominated interest rate implied from an foreign exchange swap. The Russian overnight interest rate swap (ROISfix) used in some specifications is available since 15 April 2011, while the Russian volatility index (RTSVX) is available since 24 January 2011. Descriptive statistics and graphs are presented in Appendix 3.

Basic MIACR is noisy due to the heterogeneity of banks in terms of risk. Later, the Central Bank of Russia introduced the MIACR-IG rate, which takes into account only interbank lending to investment-grade banks. These banks borrow less frequently from the Russian money market. As a result the number of deals per day decreasing along with increases in maturity.

To calculate MIACR-IG and implied rate spread as well as MIACR-IG and ROISfix spread we ran the following procedure: if MIACR-IG isn't observed, we mark the spread value as missing. If MIACR-IG is observed while RTSVX or ROISfix are not reported, then we drop the whole observation.

To begin the analysis, we tested the necessary series for stationarity. For Mosprime spreads, we used generic ADF tests. This test is not applicable for series with missing values as differences cannot be calculated. We therefore used the Phillips-Perron test in this case. The idea of the test is to estimate AR(1) with errors adjusted for autocorrelation in residuals. AR(1) models were estimated using a Kalman filter technique. Unit root tests for variables used in our models are presented in Table 3.

Table 3. Unit root tests[1,2]

Variable	Maturity	ADF (levels)	ADF (differences)
Mosprime – Implied USD rate spread	1 w	-7.19***	
	1 m	-4.50***	
	3 m	-2.89	-12.06***
	6 m	-2.53	-11.64***
Mosprime – ROISfix spread	1 w	-5.66***	
	1 m	-4.27***	
	3 m	-1.63	-9.69***
	6 m	-1.64	-9.02***
Fixed REPO – Implied USD rate spread	1 w	-3.06**	
	1 m	-2.55	-31.38***
	3 m	-2.02	-21.87***
	6 m	-1.74	-24.76***
ln RTSVX		-3.02**	
Variable	Maturity	PP (levels)	
MIACR-IG – Implied USD rate spread	1 w	-12.11***	
	1 m	-11.19***	
	3 m	-12.91***	
	6 m	-8.57***	
MIACR-IG – ROISfix Spread	1 w	-17.25***	
	1 m	-4.47***	
	3 m	-2.55	
	6 m	-2.34	

[1] Here and below *** refers to significance at 1% level, ** to significance at 5% level, * to significance at 10% level.

[2] McKinnon statistics are shown in the table.

As we can see, offered rate based spreads for longer maturities are I(1), while for all other variables tests reject non-stationarity hypotheses in favor of I(0).

5. Empirical strategy

5.1. Does CIP hold?

We propose to test difference between the interbank ruble rate and the FX swap implied rate for equality to zero. This way of testing CIP is the simplest one and is well suited for the analysis of deviation from parity and its causes. CIP implies a stationary deviation process with zero mean. Parity does not hold if spread is not stationary. To test for CIP in the case of stationary spread series, we estimated the AR(p) model (8).

$$s_t - \delta = \sum_{i=1}^{p} \phi_i \left(s_{t-i} - \delta\right) + \varepsilon_t, \tag{8}$$

where s_t is a spread between an interbank rate and an implied foreign interest rate, AR model order p is determined based on the Swartz Information Criterion (Shittu et al., 2009).

Null hypothesis of CIP is δ equals to zero, or more formally:

$$H_0 : \delta = 0$$
$$H_1 : \delta \neq 0$$

In most interest cases of MIACR-IG as ruble interest rates, we have a lot of missing values due to absence of deals. We propose to use the Kalman filter approach to manage this problem.

5.2. Are deviations from CIP long-lasting?

Usually, profitable opportunities of arbitrage within financial markets disappear almost immediately and would not be observable in daily data. As far as CIP does not hold on average, we may interpret temporary deviations as departures from the long-run interest spread rather than disparities. For purely transitional deviations, all coefficients of the AR model (8) are equal to zero. If the coefficients are not jointly equal to zero, however, deviations persist. We therefore test $H_0 : \phi_1 = \phi_2 = ... = \phi_p = 0$ against the natural alternative.

5.3. Are there any thresholds?

Transaction costs are possible sources of persistent deviations from parity. In this case, we may expect that small deviations will not produce any profitable arbitrage opportunities and therefore will persist. Larger deviations should, however, be short-lived. This behavior is well captured by threshold auto-regression (TAR) models. Following Hutchison et al. (2012), we constructed a three-regime model. If there are transaction costs, there would be an inner regime with small and persistent deviations from CIP and two outer regimes with larger and more persistent deviations (see Fig. 6 as an example). This behavior is unlikely when transaction costs are absent.

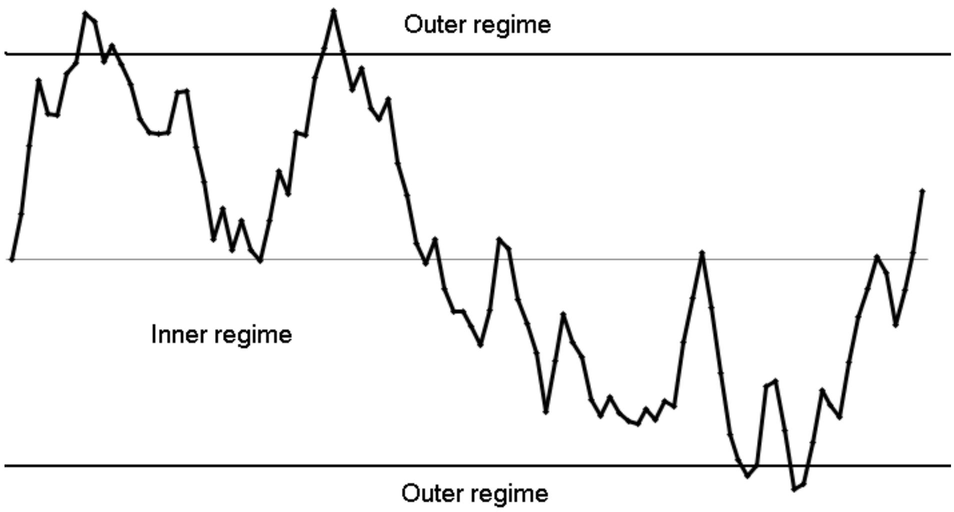

Fig. 6. Dynamics of a variable following three-regime TAR. An example.

To test for three-regime TAR we applied a procedure proposed by Hansen (1999) and following Hutchison et al. (2012).

We estimated a simple linear AR(1) model (9):

$$s_t = \delta + \phi s_{t-1} + \varepsilon_t \qquad (9)$$

Then we chose the best two-regime TAR(1) model (10):

$$s_t = \begin{cases} \delta_L + \phi_L s_{t-1} + \varepsilon_t, s_{t-1} < \tau \\ \delta_H + \phi_H s_{t-1} + \varepsilon_t, s_{t-1} > \tau \end{cases} \qquad (10)$$

Linear restrictions were tested using F-test. The distribution of the statistics is not standard, however, as the structural break was determined endogenously as a result of sequential estimation of TAR(1) models with various thresholds τ. We computed critical values based on 1000 simulations as proposed in Hansen (1999).

If there is at least one threshold, we ran the procedure of searching for the best three-regime TAR(1) model (11):

$$s_t = \begin{cases} \delta_L + \phi_{\mathrm{L}} s_{t-1} + \varepsilon_t, s_{t-1} < \tau_L \\ \delta_M + \phi_{\mathrm{M}} s_{t-1} + \varepsilon_t, \tau_L < s_{t-1} < \tau_H \\ \delta_H + \phi_{\mathrm{H}} s_{t-1} + \varepsilon_t, s_{t-1} > \tau_H \end{cases} \tag{11}$$

We used a similar procedure to compute the critical values of the test and to choose between two and three-regime models.

This test was applied to Mosprime-based spreads only as there is no appropriate procedure for time-series with missing values.

5.4. Does credit risk premium explain deviations CIP?

Counterparty risk premium is a possible explanation for the positive long-run deviations from parity. Credit default swap (CDS) premiums, the spread between risky interbank rates and virtually riskless overnight interest swap (OIS), and stock market volatility are typical proxies for the credit risk in the literature. Interbank-OIS spread is the most appropriate measure of credit risk for our study. It reflects the existing risks in the market and has the same maturities as the money rates we use. CDS premiums, on the contrary, have much longer maturities, while stock market volatility reflects stock market uncertainty and risk aversion, which indirectly relate to money market risks (Bekaert et al., 2013).

The OIS contract is associated with a relatively small counterparty risk and with very small liquidity risk premiums, as there is no principal exchange and the contract does not assume any initial cashflow (Baba and Packer, 2009a). Conversely, the interbank interest rate is far from risk-free. The interbank interest rate-ROISfix spread hence accurately reflects the credit risk in the money market.

Model (12) explains deviations from parity by the counterparty risk premium.

$$s_t = \beta_0 + \beta_1 (i_t - ROISfix_t) + \varepsilon_t \qquad (12)$$

We tested the null hypothesis that the risk premium doesn't affect deviations from CIP against the alternative of a linear relationship expecting a positive sign. Formally:

$$H_0 : \beta_1 = 0$$
$$H_1 : \beta_1 > 0$$

The Mosprime rate was applied to a generic first class potential borrower. Conversely, MIACR-IG is an average rate for investment-grade borrowers. The group of such borrowers changes from day to day depending on current deals. We therefore applied the Mosprime-ROISfix spread to capture credit risk within the Mosprime rate, while the MIACR-IG-ROISfix spread is used to explain deviations of actual rates from parity.

Appropriate econometric treatment of the model (12) depends on stationarity of explanatory and dependent time-series. If both of them were integrated of order one then the Engle-Granger procedure was applied. When the variables were stationary, we used OLS regression. Different orders of integration signal that (12) has been mis-specified. Table 4 summarizes all of the above concepts.

Table 4. Choice of estimation method

	I(0)	I(1)
I(0)	OLS	-
I(1)	-	Engle-Granger procedure

We started with Mosprime deviations from CIP. According to the unit root test (see Table 3) three and six-month maturities are not stationary. We used the two-step Engle-Granger procedure for these series. As the spreads for shorter maturity have no unit roots, we ran OLS.

The MIACR-IG implied rate spread is stationary for all maturities, while the unit root hypothesis for MIACR-IG–ROISfix spread cannot be rejected for three and six-month maturities (see Table 3). We thus applied OLS to one-week and one-month maturity data and did not estimate model (12) for longer maturities.

5.5. Does turbulence explain deviations from CIP?

There is some evidence that deviations from CIP are associated with turbulence in financial markets. Branson (1969) and Taylor (1989) note that deviations from parity grow with turbulence. Skinner and Mason (2011) also attributed deviation from CIP in the Russian money market to instability in financial markets.

Stock market volatility index is a natural measure of turbulence in financial markets. RTSVX is a measure of expected volatility in the Russian stock market, which is similar to VIX for the US market. We tested the impact of RTSVX on deviations from CIP (13).

$$\mathrm{s_t} = \beta_0 + \beta_1 \ln \mathrm{RTSVX}_{\ \mathrm{t}} + \varepsilon_t \qquad (13)$$

We can formulate our null hypothesis as $H_0 : \beta_1 = 0,$ while the alternative is $H_1 : \beta_1 > 0$. We did, however, expect a positive sign of β_1.

As MIACR-IG-implied rate spreads are stationary, we ran a simple linear regression of spreads on the log of RTSVX (which turned out to be stationary as well) for all maturities. One week and one month Mosprime-implied rate spreads are also stationary. We therefore used the same model. On the contrary, three and six-month Mosprime-based spreads are I(1) and cannot be explained by the stationary turbulence variable. The model must therefore be mis-specified.

Higher volatility is often accompanied by greater credit risk. This encourages us to decompose the spread into part driven by credit risk and those driven solely by turbulence. The estimation strategy differs for the stationary spreads related to shorter term spreads. Fig. 7 describes the choice of the approaches depending on properties of the time-series.

Firstly we show how to deal with stationary dependent and explanatory variables. In these circumstances there is no need to perform co-integration analysis. We suppose that while RTSVX may be associated both with credit risk and turbulence, interbank-OIS rates measures credit risk only. We therefore used regression of the spread on the residuals from model (14.1) to separate the potential sources of deviations from parity:

$$\ln \mathrm{RTSVX}_{t} = \beta_0 + \beta_1 (i_t - ROISfix_t) + u_t \tag{14.1}$$

$$\mathrm{s}_{t} = \beta_0 + \beta_1 (i_t - ROISfix_t) + \beta_2 u_t + \varepsilon_t, \tag{14.2}$$

where u_t are residuals of log of RTSVX regression on credit risk, which are orthogonal to credit risk and reflect turbulence. To determine whether turbulence affects deviations from parity, we tested the null hypothesis of no effect against an alternative of its presence: $H_0 : \beta_2 = 0; H_1 : \beta_2 \neq 0.$

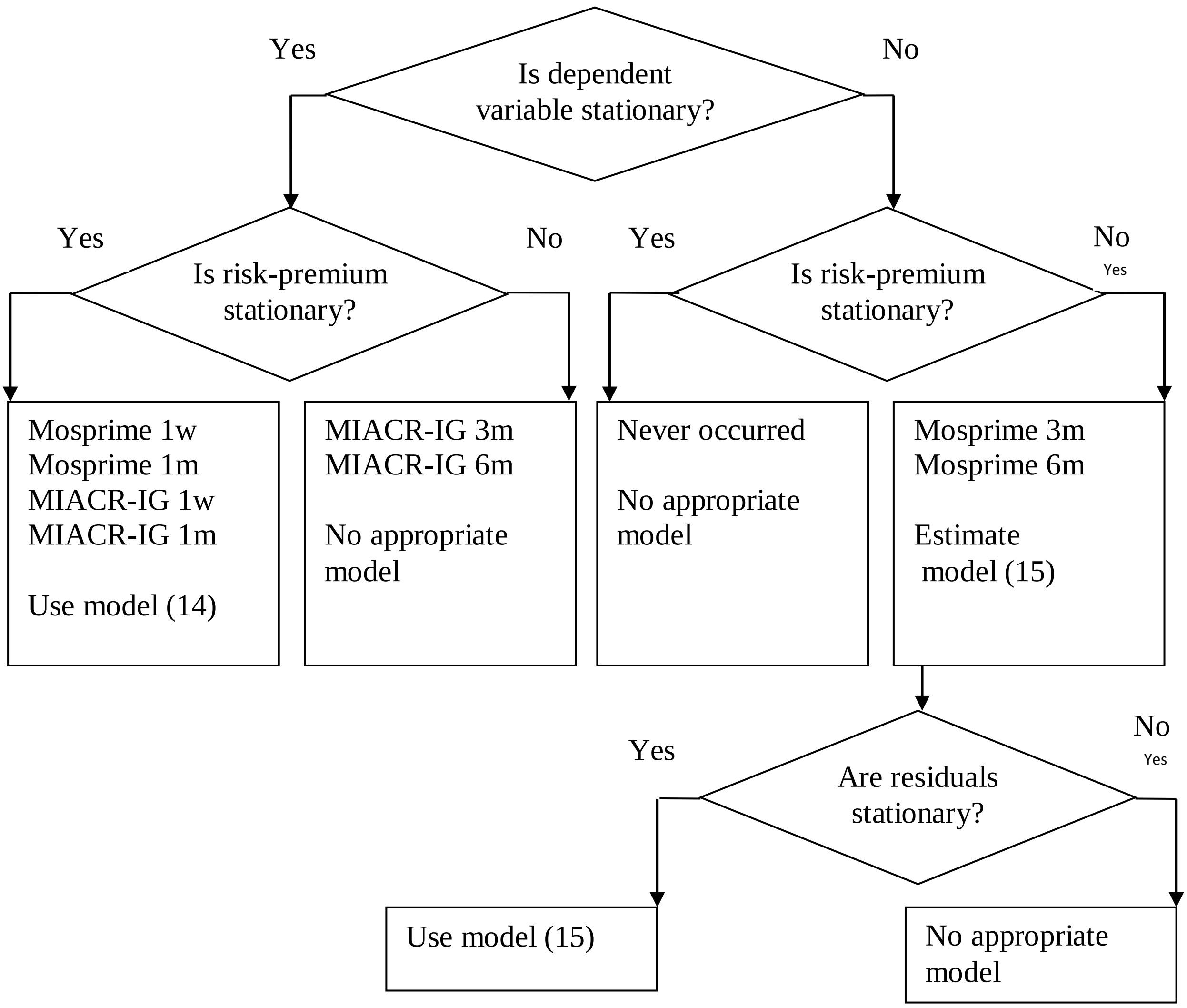

Fig. 7. Approaches to estimating both turbulence and credit risk effects

When a dependent variable is nonstationary, model (15) requires a nonstationary risk premium to avoid misspecification when the turbulence measure is I(0). We conducted the Engle-Granger procedure only for three and six-month Mosprime IG–implied rate models, which satisfy the requirement. In the case of stationary residuals, we can check if turbulence influences deviations from parity based on β_2 estimation, namely: $H_0 : \beta_2 = 0, H_1 : \beta_2 \neq 0$.

$$s_t = \beta_0 + \beta_1 (i_t - ROISfix_t) + \beta_2 \ln \mathrm{RTSVX}_t + \varepsilon_t \tag{15}$$

Finally, linear models like (14) and (15) are inapplicable to situations in which a dependent variable is stationary and only one regressor is nonstationary. We therefore did not estimate these types of models for MIACR-IG three and six-month spreads.

5.6. Does monetary policy affect CIP deviations?

The model proposed in section 3 predicts higher risk premiums when the policy rate is higher due to riskier banks turning to interbank lending. To test if this effect takes place, we regressed deviations from CIP on both our measure of risk premium and a policy rate deviation from parity (16):

$$s_t = \beta_0 + \beta_1 (i_t - ROISfix_t) + \beta_2 (\mathrm{R}EPO_t - SWAP_t) + \varepsilon_t, \tag{16}$$

where $\mathrm{R}EPO_t$ is a REPO lending standing facility rate and $SWAP_t$ is a USD implied swap rate. The latter spread is stationary only for one-week maturities (see Table 3). We thus ran a simple linear model in this case and used co-integration analysis for three and six-month maturities. The model is inapplicable to one-month data, as in this case the left hand side of the equation is stationary, while the right hand is I(1).

We tested for policy rate effect following the same way as we did in model (15). Null hypothesis of no effect implies $\beta_2 = 0$ while the alternative is $\beta_2 > 0$ as we expect a positive sign.

5.7. Estimation of risk, turbulence, and policy impact on deviations from parity

It is logical to bring all factors affecting CIP deviations together into a single model and to estimate their relative influence, as we present in (17):

$$\mathrm{s}_{\mathrm{t}} = \beta_0 + \beta_1\left(i_t - ROISfix_t\right) + \beta_2\left(\mathrm{R}EPO_t - SWAP_t\right) + \beta_3 u_t + \varepsilon_t, \qquad (17)$$

where u_t are residuals from $\ln\mathrm{RTSVX}_{\mathrm{t}}$ regression on risk premiums as in (14). Model (17) is applicable to one-week data, where all regressors and explanatory variable are stationary. In the case of three-month and six-month spreads it is necessary to check if co-integration between nonstationary variables is present. In the case of one-month data, (17) cannot be used, as left and right hand sides of the model have different orders of integration. We therefore substitute it with (15) in these circumstances.

We used the estimates to evaluate the influence of the factors on deviations from the spread. It makes sense to use two indicators. The first one is a standardized coefficient which shows for how many standard deviations the spread changes responding to one standard deviation of a particular factor. We calculate it as:

$$\hat{\beta}_i^* = \hat{\beta}_i \frac{s_i}{s_s}, \qquad (18)$$

where $\hat{\beta}_i$ is an estimation of *i*th coefficient in a model, s_i is a sample standard deviation of *i*th regressor and s_s is a sample standard deviation of a dependent variable.

Standardized coefficients, however, neglect the size of the response, which is useful for discussing sources of deviations from parity rather than deviations from the mean. We therefore reported an alternative measure, namely semi-standardized coefficients:

$$\hat{\beta}_i^+ = \hat{\beta}_i s_i, \qquad (19)$$

which shows how a one standard deviation change in *i*th factor affects CIP deviation in percentage points.

Additionally, we calculate factors' contribution to deviations from parity:

$$c_i = \frac{\hat{\beta}_i^+}{\sum_i \hat{\beta}_i^+}. \qquad (20)$$

6. Results

6.1. CIP doesn't hold on average

As we noted earlier in section 5.1, CIP doesn't hold for nonstationary spreads; that is Mosprime-based spreads for three and six-month loans. The estimation results of model (8) for stationary spreads are presented in Table 5.

We also see that CIP doesn't hold on average for stationary time-series. Average spread grows with maturity from a slightly negative value of 8 basis points for weekly MIACR-IG rate to a definitely positive 105 basis points for 6-month lending. Average deviations are higher for higher maturities. The same results are obtained in Batten et al. (2010). They found evidence that average deviation is greatest for one-year maturities and smallest for the shortest in the sample, the one-month maturity.

Table 5. Presence and persistence of deviations from CIP

Spread (s_t)	Maturity	δ[1]	p	$H_0 : \phi_1 = \phi_2 = ... = \phi_p = 0$[2]
Mosprime – Implied USD rate	1 w	0.16*** (0.02)	8	90.87***
	1 m	0.33*** (0.06)	8	31.19***
MIACR-IG – Implied USD rate	1 w	-0.08** (0.04)	5	3.82
	1 m	0.60*** (0.04)	1	239.79***
	3 m	0.44*** (0.10)	3	7.56**
	6 m	1.05*** (0.11)	3	100.70***

[1] Standard errors are in the parentheses.

[2] Test statistics is distributed as $\chi^2(p)$.

6.2. Deviations from CIP are mostly persistent

As can be seen in Table 5, autoregressive coefficients in model (8) are not jointly equal to zero for all models except one. This demonstrates clearly that the deviations are not purely transitory for all maturities except for the weekly one for the MIACR-IG spread. These results raise a question about causes of CIP violation.

6.3. Transaction costs are small

We present here the major results of TAR estimations in Table 6. Simulated critical values are in Table 7.

Table 6. Testing for thresholds

	One threshold		Two thresholds	
	Upper regime	F-statistics	Inner regime	F-statistics
1 w	[0.58;∞)	24.31***	[0.05; 0.05]	19.98***
1 m	[0.17;∞)	8.81***	[-0.24; 0]	3.27*
3 m	[0.27;∞)	16.04***	[-0.12; 0.06]	4.85***
6 m	[-0.07;∞)	18.91***	[-0.07; 0.03]	1.55

Table 7. Critical values for F-test (based on 1000 simulations)

	One threshold vs No thresholds			Two thresholds vs One threshold		
	0.1	0.05	0.01	0.1	0.05	0.01
1 w	5.08	5.63	7.53	3.09	3.82	5.69
1 m	5.21	6.25	7.39	2.77	3.34	4.27
3 m	5.45	6.12	7.66	2.86	3.29	4.40
6 m	5.63	6.29	8.16	3.02	3.62	5.07

Tests rejected the hypothesis of linearity for all the Mosprime spread series. The estimated neutral band is narrow, indicating low transaction costs. TAR models, however, strongly favor two thresholds for the series related to one-week and three-month maturities only. For one-month maturity, the threshold presence is marginally significant. The estimation results are not robust, however. Moderate changes in estimation techniques significantly affect the bandwidth and estimates of autoregression coefficients.

6.4. Credit risk drives the spread

The estimation results for the credit risk model (12) are presented in Table 8. We can see that credit risk is highly significant for all maturities. The impact of credit risk sharpens as maturities grow. One percentage point growth in a six-month risk premium adds nearly the same value to deviation from parity. Risk premium effects on actual rate spreads are stronger compared to those based on offered rates for the same maturities.

Table 8. Risk premium regressions

Model	Maturity	β_0	β_1	ADF (levels)
Dependent variable is MosPrime – Implied USD rate spread				
$s_t = \beta_0 + \beta_1(i_t - ROISfix_t) + \varepsilon_t$	1w	0.02 (0.02)	0.25*** (0.05)	
	1m	-0.02 (0.02)	0.50*** (0.04)	
$s_t = \beta_0 + \beta_1(i_t - ROISfix_t) + \varepsilon_t$	3m	-0.20*** (0.03)	0.80*** (0.03)	-4.18***
	6m	-0.34*** (0.03)	0.98*** (0.03)	-3.30**
Dependent variable is MIACR-IG – Implied USD rate spread				
$s_t = \beta_0 + \beta_1(i_t - ROISfix_t) + \varepsilon_t$	1w	-0.21*** (0.02)	0.60*** (0.06)	
	1m	-0.30*** (0.03)	0.88*** (0.04)	

6.5. Turbulence tends to diminish the spread

Table 9 presents the estimation results for model (13). Volatility in the Russian stock market does not significantly impact the spread. Similar results were obtained by Skinner and Mason (2011), who explained deviation from CIP for Brazil, Chile, Russia, South Korea, Norway and the UK. They found an ambiguous influence of the VIX variable. It has significantly positive effect only for the OLS model estimated on five-year data for Brazil and insignificantly or even negatively impacts the spread for the rest datasets.

Table 9. Turbulence regressions

Model	Maturity	β_0	β_1
Dependent variable is MosPrime – Implied USD rate spread			
$s_t = \beta_0 + \beta_1 \ln RTSVX_t + \varepsilon_t$	1w	-0.04 (0.10)	0.05** (0.03)
	1m	0.29*** (0.16)	0.00 (0.05)
Dependent variable is MIACR-IG – Implied USD rate spread			
$s_t = \beta_0 + \beta_1 \ln RTSVX_t + \varepsilon_t$	1w	-0.39 (0.28)	0.06 (0.08)
	1m	-0.09 (0.48)	0.09 (0.14)
	3m	1.62** (0.64)	-0.31 (0.19)
	6m	1.44 (0.97)	-0.20 (0.28)

Stock market volatility and credit risk interaction might produce this ambiguity. We decomposed these influences, as discussed in Section 5.5, presenting estimation results for models (14) and (15) in Table 10. The turbulence component is significantly negative now in all models except for the weekly Mosprime spread.

Table 10. Risk and turbulence regressions

Model	Maturity	β_0	β_1	β_2	ADF (levels)
Dependent variable is MosPrime – Implied USD rate spread					
$\ln\mathrm{RTSVX}_t = \beta_0 + \beta_1 (i_t - ROISfix_t) + u_t$	1w	3.19*** (0.03)	0.50*** (0.07)		
	1m	3.00*** (0.03)	0.56*** (0.04)		
$\mathrm{s}_t = \beta_0 + \beta_1 (i_t - ROISfix_t) + \beta_2 u_t + \varepsilon_t$	1w	0.02 (0.03)	0.25*** (0.08)	-0.00 (0.03)	
	1m	-0.02 (0.03)	0.50*** (0.04)	-0.24*** (0.03)	
$\mathrm{s}_t = \beta_0 + \beta_1 (i_t - ROISfix_t) + {} + \beta_2 \ln\mathrm{RTSVX}_t + \varepsilon_t$; $\varepsilon_t - \varepsilon_{t-1} = \beta_3 \varepsilon_{t-1} + \xi_t$	3m	1.62*** (0.10)	0.85*** (0.03)	-0.56*** (0.03)	-5.11***
	6m	1.36*** (0.10)	0.97*** (0.03)	-0.50*** (0.03)	-5.82***
Dependent variable is MIACR-IG – Implied USD rate spread					
$\ln\mathrm{RTSVX}_t = \beta_0 + \beta_1 (i_t - ROISfix_t) + u_t$	1w	3.30*** (0.02)	0.21*** (0.06)		
	1m	3.29*** (0.03)	0.17*** (0.04)		
$\mathrm{s}_t = \beta_0 + \beta_1 (i_t - ROISfix_t) + \beta_2 u_t + \varepsilon_t$	1w	-0.21*** (0.06)	0.60*** (0.07)	-0.10*** (0.07)	
	1m	-0.30*** (0.03)	0.88*** (0.04)	-0.43*** (0.07)	

6.6. Policy affects the spread

Table 11 shows estimation results for the policy rate model (16). As we can see, the spread between the policy rate and implied rate increases domestic money market rates compared to foreign ones. This holds true both for Mosprime and for MIACR-IG rates. Policy effect is stronger for longer maturities.

Table 11. Policy rate regressions

Model	Maturity	β_0	β_1	β_2	ADF (levels)
Dependent variable is MosPrime – Implied USD rate spread					
$s_t = \beta_0 + \beta_1(i_t - ROISfix_t) + \beta_2(REPO_t - SWAP_t) + \varepsilon_t$	1w	-0.10*** (0.02)	0.39** (0.05)	0.09*** (0.01)	
	3m	-0.57*** (0.04)	1.07*** (0.03)	0.18*** (0.01)	-4.05***
	6m	-0.80*** (0.04)	1.27*** (0.03)	0.22*** (0.01)	-4.63***
Dependent variable is MIACR-IG – Implied USD rate spread					
$s_t = \beta_0 + \beta_1(i_t - ROISfix_t) + \beta_2(REPO_t - SWAP_t) + \varepsilon_t$	1w	-0.31*** (0.02)	0.64*** (0.06)	0.14*** (0.02)	
	3m	-0.48*** (0.03)	1.00*** (0.02)	0.16*** (0.02)	-3.72***
	6m	-0.33*** (0.08)	1.00*** (0.01)	0.59*** (0.05)	-2.75**

6.7. Risk premium, turbulence, and policy matter

Table 12 demonstrates the estimations of model (17) which includes all analyzed factors.

Table 12. Long model

Model	Maturity	β_0	β_1	β_2	β_3	ADF (levels)
Dependent variable is MosPrime – Implied USD rate spread						
$\ln RTSVX_t = \beta_0 + \beta_1(i_t - ROISfix_t) + u_t$	1w	3,19*** (0.02)	0.5*** (0.07)			
$s_t = \beta_0 + \beta_1(i_t - ROISfix_t) + \beta_2(REPO_t - SWAP_t) + \beta_3 u_t + \varepsilon_t$	1w	-0.11** (0.05)	0.41*** (0.09)	0.10*** (0.02)	-0.11** (0.05)	
$s_t = \beta_0 + \beta_1(i_t - ROISfix_t) + \beta_2(REPO_t - SWAP_t) + \beta_3 \ln RTSVX + \varepsilon_t$	3m	1.29*** (0.1)	1.14*** (0.03)	0.19*** (0.01)	-0.58*** (0.03)	-5.47***
	6m	0.77*** (0.09)	1.24*** (0.02)	0.20*** (0.01)	-0.46*** (0.03)	-6.19***
Dependent variable is MIACR-IG – Implied USD rate spread						
$\ln RTSVX_t = \beta_0 + \beta_1(i_t - ROISfix_t) + u_t$	1w	3.30*** (0.02)	0.21*** (0.06)			
$s_t = \beta_0 + \beta_1(i_t - ROISfix_t) + \beta_2(REPO_t - SWAP_t) + \beta_3 u_t + \varepsilon_t$	1w	-0.11** (0.05)	0.64*** (0.06)	0.15*** (0.02)	-0.19** (0.06)	
$s_t = \beta_0 + \beta_1(i_t - ROISfix_t) + \beta_2(REPO_t - SWAP_t) + \beta_3 \ln RTSVX + \varepsilon_t$	3m	1.48*** (0.16)	1.02*** (0.02)	0.15*** (0.01)	-0.58*** (0.05)	-5.59***
	6m	1.33*** (0.24)	1.00*** (0.02)	0.12*** (0.02)	-0.51*** (0.07)	-3.72***

Noticeably, all factors included in the model are significant. Qualitative results described in sections 6.4-6.6. persist. Credit risk effect is strongly positive, especially for longer maturities. Turbulence has a significantly negative effect for all spreads, including weekly Mosprime. Finally, the spread between the fixed REPO rate and implied swap rate shifts deviations from CIP upwards. Quantitatively, estimates are quite close to those obtained previously.

Table 13. Factors' contribution to the spread

Dependent variable	MosPrime – Implied USD rate spread		MIACR-IG – Implied USD rate spread	
Maturity	1w	1m	1w	1m
	Standardized coefficients			
Risk premium	0.35	0.54	0.56	0.87
Turbulence	-0.15	-0.28	-0.15	-0.23
Monetary policy	0.50		0.50	
	Semistandardized coefficients			
Risk premium	0.07	0.12	0.19	0.51
Turbulence	-0.03	-0.06	-0.05	-0.13
Monetary policy	0.11		0.16	
	Per cent contribution to deviation from parity			
Risk premium	50%	208%	62%	135%
Turbulence	-22%	-108%	-16%	-35%
Monetary policy	72%		55%	

Table 13 shows the influences of the discussed drivers on deviations from parity for stationary series models. Risk premium and monetary policy are major forces affecting the spread, while the influence of turbulence is smaller and negative. Risk premium growth, for example, for one standard deviation raises the weekly Mosprime-implied rate spread rate by 0.35 standard deviations or 7 basis points. This is approximately half of the average spread, which is equal to 17 basis points.

7. Conclusion

This study has examined the presence, scale and causes of deviations from covered interest parity (CIP) using actual and offered rates in the Russian money market between 2010-2014.

Baseline results exhibited CIP violations. Internal rates exceeded international money market rates after accounting for forward premiums. This effect is more pronounced for longer maturities and for actual rates compared to the offered ones. For example, the MIACR-IG rate on monthly lending is 60 basis points higher than the implied USD rate and goes 27 points on average beyond the corresponding Mosprime rate. The MIACR-IG rate on six-month loans exceeds the implied rate for 105 basis points. Moreover, deviations of money market–implied rate spreads from their average values are quite persistent. A notable exception is the weekly MIACR-IG spread, which is as low as -8 basis points with purely transitory deviations.

We tested three explanations for the CIP violation, namely credit risk, turbulence on financial markets, and policy rate effect.

Our findings indicate that credit risk affects the spread significantly. Two features are noticeable. Firstly, risks on longer borrowings are more volatile, while spreads on these borrowings are more responsive to these risks. Secondly, exactly the same holds true for the actual rates in comparison to Mosprime. The role of credit risk is therefore crucial for long-term interbank lending and for actual conditions in money markets (in contrast to those reflected by offered rates). One standard deviation of risk premium, for example, adds 12 basis points to one month Mosprime–implied rate spread, while this effect grows up to 51 points for the MIACR-IG spread, both due to higher variance of risk premiums and the stronger impact of these on deviations from parity. Baba and Packer (2009a) highlight that risk premiums had a significant effect on deviations from parity for three-month contracts in developed markets during the last financial crisis and reported similar estimates.

Several studies emphasize that deviations from parity are greater during periods of turbulence. Attempts to explain deviations from parity by VIX, however, were unsuccessful (Skinner and Mason, 2011). This might occur due to the fact that credit risk typically goes together with turbulence. To avoid this effect, we isolated part of the Russian volatility index which is orthogonal to credit risk and noticed its significantly negative impact on the spreads. This effect is moderate. 10% growth in the volatility index decreases the spread by one to four basis points provided that credit risk stays the same. This effect is stronger for longer maturities and actual rates.

The marginal effect of monetary policy is quite small and homogeneous across the rates and maturities. The positive sign matches our predictions. Spread between policy and implied rates vary significantly with time, however. One standard deviation of the spread therefore has a sizeable effect on deviations from CIP. It adds 11 basis points and 16 basis points for Mosprime and MIACR-IG-implied USD rate spreads respectively for weekly loans.

Our study concludes that the evidence from actual and offered rates is mostly the same. Credit risk and turbulence effects are more pronounced for MIACR-IG spreads than for Mosprime.

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Appendices

Appendix 1. Abbreviations.

ADF	augmented Dickey–Fuller
AR model	autoregressive model
b.p.	basis points
CBR	the Central Bank of the Russian Federation
CDS	credit default swap
CIP	covered interest parity
EURIBOR	Euro interbank offered rate.
FX	foreign exchange
GARCH model	generalized dutoregressive conditional heteroscedasticity model
I(0)	integration of order zero
I(1)	integration of order one
LIBOR	London interbank offered rate
MIACR	Moscow interbank actual credit rate. Calculated by CBR.
MIACR-IG	Moscow interbank actual credit rate- investment grade. Calculated by CBR
MIBID	Moscow interbank bid. Calculated by CBR
MIBOR	Moscow interbank offered rate. Calculated by CBR
Mosibor	Moscow interbank offered rate. Calculated by NFEA in 2001-2010.
Mosprime	Moscow prime offered rate. Calculated by NFEA
NFEA	National foreign exchange association
OIS	overnight interest swap
PP	Phillips-Perron
p.p.	percentage point
REPO	repurchase agreement
ROISfix	RUONIA Overnight Interest Rate Swap
RUONIA	Rouble overnight index average
RTSVX	Russian volatility of
TAR	threshold autoregressive model
TED spread	difference between the 3 month Treasury bill rate and the 3 month LIBOR
OLS	ordinary least squares
USD	U.S. Dollar
VIX index	implied volatility of S&P 500 index options
1w	one week
1m	one month
3m	three months
6m	six months

Appendix 2. Monetary policy and deviations from CIP. Derivations.

Let's derive the supply of and demand for credit resources as functions of a risk-free rate. Supply is equal to share of lending banks multiplied by *h*. No banks ready to provide loans at money market if risk-free rate is below the deposit rate.

$$L_s = \begin{cases} h\frac{i}{\bar{i}}, i \geq d \\ 0, i < d \end{cases}. \tag{A1}$$

As it can be seen in Fig. 6, the demand depends on whether the central bank lending rate is higher than the reservation price for liquidity or not:

$$L_d = \begin{cases} a\left[\frac{(r_x - i)^2 + 2(\bar{i} - r_x)(r_x - i)}{2\overline{rp} \cdot \bar{i}}\right], i \leq \mathrm{r_x} < \bar{i} \\ a\frac{(\bar{i} - i)^2}{2\overline{rp} \cdot \bar{i}}, r_x \geq \bar{i} \end{cases}. \tag{A2}$$

The equilibrium risk-free rate balances demand and supply. Solving quadratic equations for two cases and taking intervals into account obtains (A3):

$$i^* = \begin{cases} \bar{i} + \overline{rp}\frac{h}{a} - \sqrt{\left(\bar{i} + \overline{rp}\frac{h}{a}\right)^2 - r_x(2\bar{i} - r_x)}, i \leq \mathrm{r_x} < \bar{i} \\ \bar{i} + \overline{rp}\frac{h}{a} - \sqrt{\overline{rp}\frac{h}{a}\left(2\bar{i} + \overline{rp}\frac{h}{a}\right)}, r_x \geq \bar{i} \end{cases}. \tag{A3}$$

To calculate the average risk-premium we integrated risk-premium over borrowing banks. These banks constitute a triangle when $r_x \geq \bar{i}$, and $\tilde{r}p = \frac{\int_0^{\bar{i}-i^*}(\bar{i} - i^* - rp)rpdrp}{\frac{(\bar{i} - i^*)^2}{2}}$. The set of borrowing banks turns into a right trapezoid if $i \leq r_x < \bar{i}$, and $\tilde{r}p = \frac{\int_0^{r_x - i^*}(\bar{i} - i^* - rp)rpdrp}{\frac{(r_x - i^*)^2}{2} + (\bar{i} - r_x)(r_x - i^*)}$. After integration and rearrangements we have:

$$\tilde{r}p = \begin{cases} \frac{r_x - i^*}{3}\left(1 + \frac{\bar{i} - r_x}{(\bar{i} - i^*) + (\bar{i} - r_x)}\right), i \leq r_x < \bar{i} \\ \frac{1}{3}(\bar{i} - i^*), r_x \geq \bar{i} \end{cases}. \tag{A4}$$

Proposition 1. Average risk-premium grows as lending rate r_x increases iff $i \leq r_x < \bar{i}$.

Proof. From (A3) and (A4) is obvious, that average risk premium does not depend on r_x if $r_x \geq \bar{i}$. Average risk premium depends both on r_x and i^* from (A4) and equilibrium risk-free interest rate i^* depends on r_x if $r_x < \bar{i}$. Let's focus on the latter case. We differentiate $\tilde{r}p$ totally in respect to r_x. After rearrangements we have:

$$\frac{d\tilde{r}p}{dr_x} = \frac{\left(\bar{i} - r_x\right)}{3\left(2\bar{i} - i^* - r_x\right)^2 \left(\bar{i} - i^* + \overline{rp}\frac{h}{a}\right)} \left(2\left(\left(\bar{i} - r_x\right) + 2\left(\bar{i} - i^*\right)\right)\overline{rp}\frac{h}{a} + \left(r_x - i^*\right)^2\right) > 0, \tag{A5}$$

Q.E.D.

Appendix 3. Descriptive statistics and graphics of spreads

Table A1

Descriptive statistics of MIACR-IG – Implied USD Rate Spreads (1w, 1m, 3m, 6m)

	1w	1m	3m	6m
Mean	-0.086431	0.198947	0.515986	0.874962
Median	-0.08	0.13	0.345	0.79
Maximum	2.1	2.24	4.3	4.48
Minimum	-1.56	-2.31	-2.63	-3.47
Std. Dev.	0.338193	0.556795	0.792114	0.9826
Observations	311	209	284	131

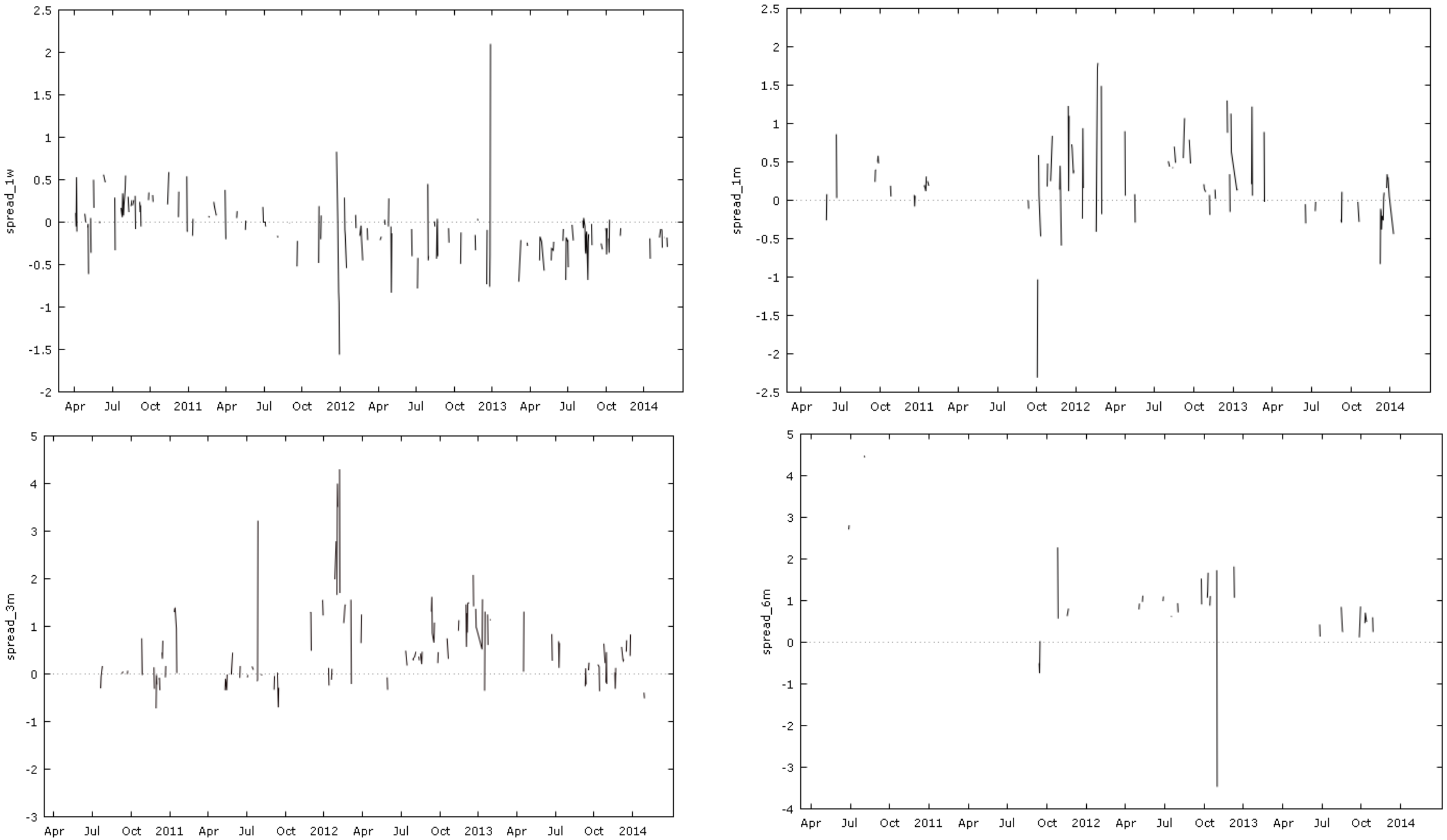

Fig.A1. MIACR-IG – Implied USD Rate Spreads (1w, 1m, 3m, 6m)

Table A2
Descriptive statistics of Mosprime – Implied USD Rate Spreads (1w, 1m, 3m, 6m)

	1w	1m	3m	6m
Mean	0.168433	0.3481	0.618225	0.752194
Median	0.15	0.33	0.61	0.75
Maximum	2.39	1.34	1.56	1.81
Minimum	-2.19	-1.49	-1.21	-0.98
Std. Dev.	0.243879	0.300701	0.388368	0.457223
Observations	958	958	958	958

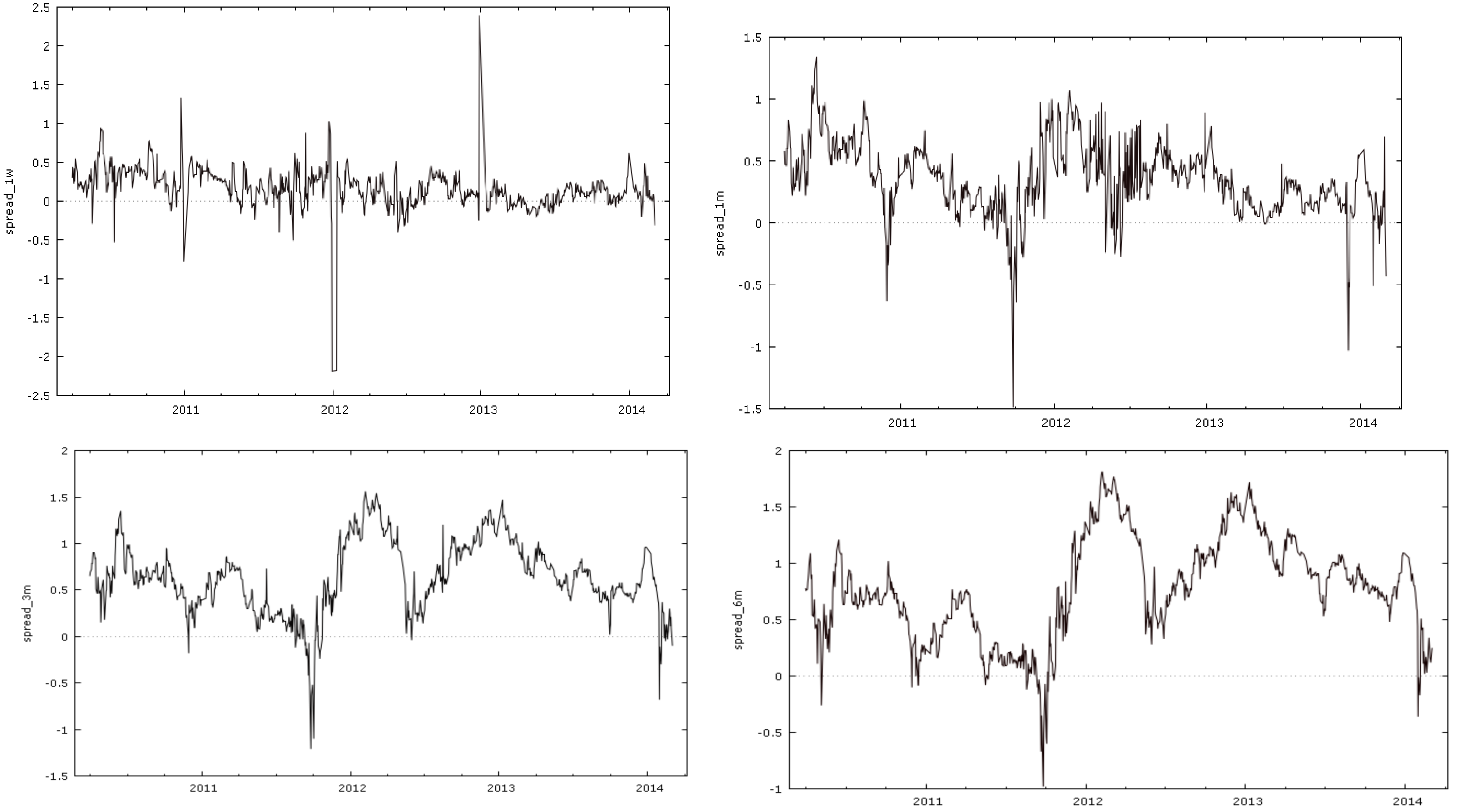

Fig.A2.Mosprime-ImpliedUSDRate Spreads (1w, 1m, 3m, 6m)

Table A3

Descriptive statistics of Miacr-IG – ROISfix Spreads (1w, 1m, 3m, 6m)

	1w	1m	3m	6m
Mean	0.048368	0.567805	0.978407	1.085377
Median	0.03	0.435	0.86	1.1
Maximum	1.22	2.95	4.41	3.04
Minimum	-0.86	-0.9	-1.33	-3.19
Std. Dev.	0.290559	0.589888	0.817924	0.828355
Observations	239	164	226	106

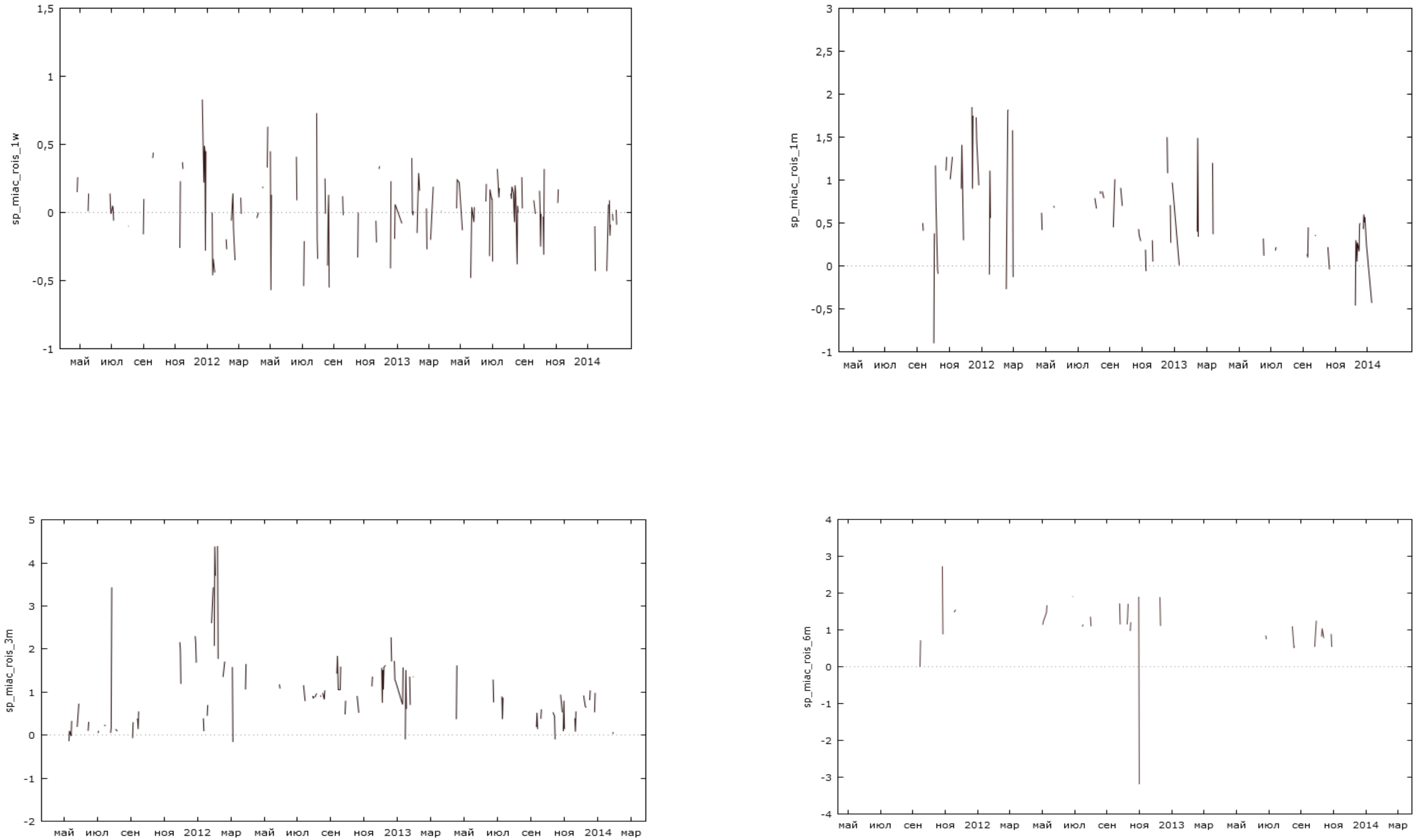

Fig.A3. Miacr-IG – ROISfix Spreads (1w, 1m, 3m, 6m)

Table A4

Descriptive statistics of Mosprime – ROISfix Spreads (1w, 1m, 3m, 6m)

	1w	1m	3m	6m
Mean	0.354197	0.340685	1.02363	1.173181
Median	0.34	0.29	1.1	1.22
Maximum	1.18	2.18	1.91	1.94
Minimum	-0.47	-0.28	0.18	-0.06
Observations	686	686	686	686

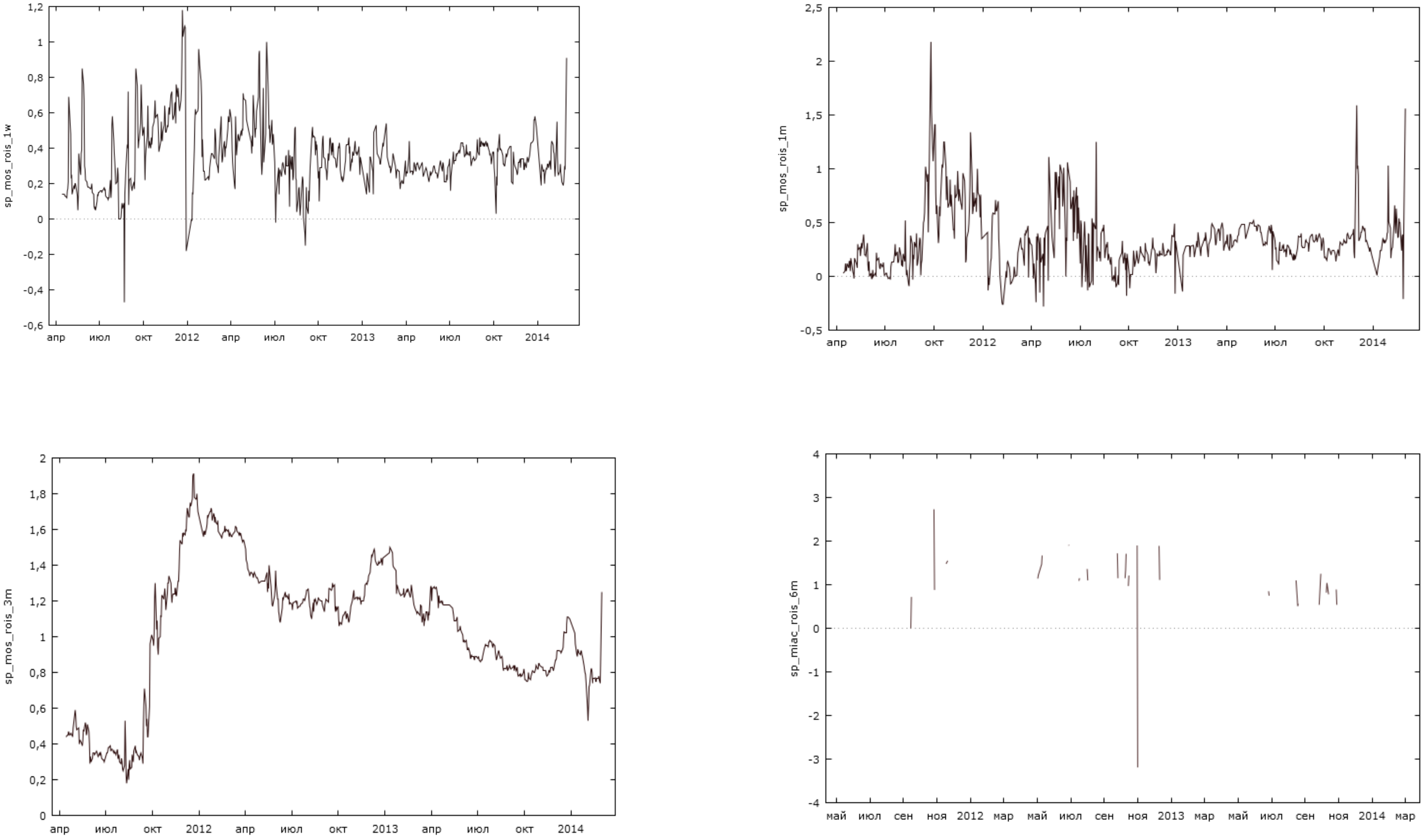

Fig.A4. Mosprime – ROISfix Spreads (1w, 1m, 3m, 6m)

Table A5

Descriptive statistics of ln(RTSVX)

Mean	3.364918
Median	3.291643
Maximum	4.310262
Minimum	2.735665
Std. Dev.	0.301624
Observations	811

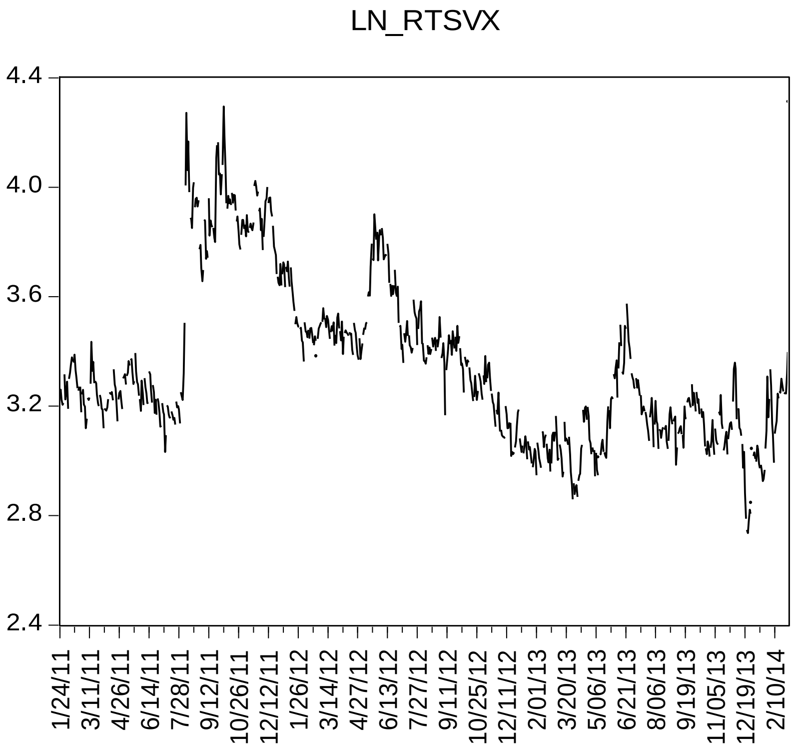

Fig.A5. ln(RTSVX) (1w, 1m, 3m, 6m)