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BASIS RISK AND NET INTEREST INCOME OF BANKS

Piotr Mielus, Ph.D.

Warsaw School of Economics, Collegium of Economic Analysis, Department of Quantitative Economics
Madalińskiego 6/8, 02-513 Warszawa, Poland
e-mail: piotr.mielus@sgh.waw.pl

Tomasz Mironczuk, MSc

Gdańsk Institute for Market Economics, Warsaw Branch
Kołobrzeska 16, 02-923 Warszawa, Poland
e-mail: tomasz.mironczuk@ibngr.pl

Anna Zamojska, Ph.D., Associate Prof.

University of Gdańsk, Faculty of Management, Department of Econometrics
Armii Krajowej 101, 81-824 Sopot, Poland
e-mail: anna.zamojska@ug.edu.pl

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Abstract

The results of the banking sector are shaped primarily by commissions and net interest income. Net interest income is determined by the difference between the profitability of bank assets and liabilities. In the case when a different method is used to determine interest for each side of the balance sheet, there occurs a basis risk that may lead to the deterioration in the net interest income of the sector. This is the situation in the Polish banking sector, which is characterized by the presence of variable interest rates for long-term assets and fixed interest rates for short-term liabilities. The study aims to verify the following thesis: in an environment of falling interest rates we can observe the deterioration in net interest income of the banking sector, as a result of the materialization of the basis risk. The authors of the article state that the source of the basis risk is the mismatch between the reference rate used to define the interest flow of loans and the actual cost of financing the balance through term deposits collected from non-financial entities.

Keywords: net interest income of the bank, the cost of financing, interest rates on deposits, financial market indices

JEL classification: G21, G32

Introduction

In 2014 the Polish banking sector generated a net profit of PLN 16 billion. The main revenue component for the banks was the result on banking activity in the amount of PLN 57.7 billion, of which 64.4% was income from interest.[1] Net interest income, after taking into account the cost of liabilities, is therefore a key source of profit in the banking sector. Analysing the structure of this result, we should consider the profitability of both sides of the balance sheet: assets represented to a major extent by loans, and liabilities generated largely by deposits.

Banks finance long-term assets with short-term liabilities, which are characterized by varying degrees of stability. Deposits from private individuals are characterized by the highest stability, due to a high tendency to extend them and a low level of volatility of the so-called residues sums maintained in current accounts. On the other hand, deposits of enterprises and the public sector are less stable, due to the higher volatility of the liquidity situation of enterprises and public entities, as well as price competition in the deposit market. Deposits made by professional entities, namely non-banking financial institutions and banks, on the interbank market, are the least stable. Due to regulatory requirements CRD IV (LCR and NSFR),[2] financial stability is taken into account in the calculation of capital adequacy, which makes deposits from individuals valued at a higher level than other deposits (i.e. their yields are usually higher than the interest rate on deposits received from other entities) (see Ötker-Robe, Pazarbasioglu, et al., 2010).

In Polish conditions, short-term deposits are dominant (with maturity of less than one year), on which interest is paid on the maturity date of the deposit and determined on the basis of a fixed rate. On the other hand, long-term loans are predominantly indexed to a variable rate based on the WIBOR[3] plus bank margin. Net interest income is thus determined by the difference between the average cost of liabilities (mainly defined by the current cost of fixed deposits) and the profitability of assets (mainly determined by the current level of WIBOR modified by a fixed margin determined on the day the loan is granted). This result will be stable in subsequent periods (with a constant volume of deposit and loans streams), provided there is a correlation between interest rates on the deposits and the WIBOR.

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[1] The data of the Financial Supervision Authority on the following website: http://www.knf.gov.pl/opracowania/sektor_bankowy/dane_o_rynku/Dane_miesieczne.html, table 06.

[2] This regards the EU directive on capital requirements, introducing (based on the guidelines of the Basel Committee on Banking Supervision, the so-called Basel III of 2010) the requirement regarding key indicators of liquidity (Liquidity Coverage Ratio and Net Stable Funding Ratio).

[3] WIBOR – Warsaw Interbank Offered Rate – a reference index showing the average cost of the offered interbank deposits based on declarations made by panelist banks at 11:00 am.

The purpose of the study is to verify the degree of the above mentioned correlation and to identify the factors observed in the money market and the effects on net interest income. Determinants of the net interest income include, on the one hand, the current level of WIBOR for periods used in the indexation of loans (1, 3 and 6 months), and on the other hand, the interest rate on total deposits reported by NBP[4] and deposits negotiated according to SMRP,[5] which are a benchmark for the marginal cost of funding weighted by its denomination, reported by the banks on the basis of the concluded transactions.

The analysis of the factors affecting net interest income of the banking sector is part of the discussion which started in a period of high volatility of interest rates after the inflationary impact of the second oil crisis. Ho and Saunders (1981) drew attention to the lack of simultaneity in the revaluation of assets and liabilities of banks in an environment of rising interest rates, which resulted in a deterioration of the net interest margin of financial institutions. On the other hand, Demirgüç-Kunt and Huizing (1999) pointed out the impact of outstanding loans on the declining net interest income, despite the theoretically positive margin generated by the nominal differences in yields on assets and liabilities. However, in the case of a low share of the so-called NPL (non-performing loans), this factor is negligible. The second factor susceptible to distort the net interest income are taxes, both those paid by banks, as well as those collected as withholding tax levied from depositories. Usually, however, the net tax effect is insignificant due to similar methods of taxation of both sides of the balance sheet.

English (2002) raised the question of the sources of volatility in net interest income, paying attention to the different bases adopted in customer contracts for assets and liabilities and the discrepancies in interest repricing dates for both sides of the balance sheet. English's analysis concerned major developed economies, while Peng (2003) focused on the banking market of small open economies with the example of Hong Kong, pointing to the risk resulting from the varied dynamics of interest rates denominated in different currencies and different sensitivities of interest rates relating to loans and deposits. A similar analysis was conducted by Maes (2004) for the Belgian market. Taking into consideration the characteristics of the financial market in Western Europe, the author draws attention to the risk of financing long-term fixed-coupon assets with short-term liabilities, the interest of which is refixed depending on the market situation. In order to reduce interest rate risk, banks enter into Interest Rate Swap

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[4] Statistics of interest rates of the National Bank of Poland: http://www.nbp.pl/home.aspx?f=/statystyka/pieniezna_i_bankowa/oprocentowanie_n.html.

[5] The Money Market Monitoring System, organized by the Gdańsk Institute for Market Economics, collects data on the prices of concluded transactions concerning fixed-term deposits negotiated in PLN and received by domestic banks, divided into 7 maturity periods and 5 customer segments (data available for participants in the System).

(IRS) contracts, which allow them to convert exposure from a fixed rate to a variable rate. The relationship between a variable rate (equal to the current IBOR rate adequate for a given currency) and the cost of liabilities may vary, which means the materialization of the basis risk.

It is worth noting that the above situation is different from the Polish case, due to the lack of significant fixed-coupon assets in the balance sheets of Polish banks. The basis risk on the Polish market results from the non-parallel change in the profitability of long-term floating rate assets indexed to WIBOR and the cost of short-term fixed-coupon liabilities. However, as Brousseau, Chailloux and Durré demonstrate (2013), the basis risk also exists on a market dominated by loans granted at a fixed rate, and there is a basis risk, as the available derivative hedging instruments against interest rate risk are indexed to EURIBOR or LIBOR, which are not correlated with the actual cost of liabilities.

1. Empirical study

In order to investigate the factors influencing the materialization of the basis risk described above, the net interest income reported in the Polish banking sector was analysed. Components of the above-mentioned result, namely measures of profitability of both sides of the balance sheet, were selected as explained variables, whereas the WIBOR reference index and measures of deposit prices relating both to total deposits, as well as negotiated deposits, were selected as explanatory variables.

1.1. Analysis of time-series properties

The starting point of the empirical study was the analysis of time-series properties of the variables in the designed models: the profitability of loans by the Financial Supervision Authority (KNF), calculated as the annualized quotient of interest income and the volume of assets according to the KNF report for a given month (RK_t), the rate on deposits according to KNF, calculated as the annualized quotient of interest expense and the volume of liabilities according to the KNF report for a given month (RD_t), monthly margin by KNF, calculated as the difference between the profitability of loans and deposits as reported by KNF (M_t), the average WIBOR, calculated as the arithmetic mean of the periods 1M, 3M and 6M for a given month (WB_t), the average rate on deposits by the National Bank of Poland (NBP), the average interest rate on new term deposits in total reported by banks for individual consumers and businesses for a given month (SD_t), the average monthly rate reported by SMRP (the cost of negotiated deposits received by banks on the basis of concluded transactions) calculated as the arithmetic mean of the 1M, 3M and 6M rates (SS_t), spread by SMRP, calculated as the difference between the SMRP rate and WIBOR rate for the month (SP_t). The study covers the period from January 2010 to July 2015 (monthly data)[6]and apart from the variables listed below, it also includes two dummy variables: $D11_t$ – the neutralization of the effect of the conversion of Polbank EFG into Polbank SA,[7] $D13_t$ – the neutralization of the effect of changes in accounting policies from December 2013 in the amount of – 1 p.p.[8]

It is worth noting that the margin according to KNF indicates the difference between income from assets and the cost of liabilities normalized to the size of the interest rate on an annual basis. On the other hand, the spread by SMRP indicates the deviation of the marginal cost of liabilities (represented by the prices of negotiated deposits) from the reference index which is the basis for the valuation of assets (directly – as in floating-coupon loans and indirectly – as a result of the collateral of the balance sheet through Interest Rate Swap contracts indexed to WIBOR).

The stationarity of the time-series of variables was tested by the augmented Dickey-Fuller test (ADF), whereas the cointegration of pairs of variables was tested by means of the Engle-Granger test (EG).The results of the ADF test on the presence of a unit root in the time-series for the value of variables confirmed the existence of the unit root in the time-series of the considered variables, and thus the time-series of all variables are integrated to the first order. The applied Engle-Granger test has shown the presence of cointegration between the following time series: RK_t and WB_t, M_t and SP_t and SS_t and RD_t.

1.2. Verification of hypotheses

In the process of the empirical verification of the theses/models formulated by the authors of the designed study, the modelling of the cointegrated series using models with an error correction mechanism (ECM) or classic dynamic models ADL was applied. It was assumed that if the variables are cointegrated, there is a process of adaptation, which prevents systematic changes in the interference component in long-term relationships (Charemza, Deadman, 1997, p. 131). The models were estimated in accordance with the two-stage procedure proposed by Engle and Granger (1991). In the first stage, the long-term relationship for primary variables was estimated and subsequently, in the second stage, the lagged deviation from the long term

[6] Data available for SMRP covers data from November 2012.

[7] The transformation of Polbank EFG (branch of a Greek bank) into Polbank SA (a Polish entity) to prepare for the merger with Raiffeisen. This event had a strong one-off impact on the interest income of the sector, a conclusion which indirectly results from the NBP report.

[8] Net interest income distorted by the introduction of Recommendation U, which enforces transfers between fee income and interest income arising from bank assurance.

path was used to estimate the model of error correction mechanism. The classical method of least squares was used in the estimation process.

When the modelling dependencies of variables, whose time-series of variables are not cointegrated, the classical autoregressive distributed lag model (ADL) were used. These are models in which the dependent variable is expressed as a function of its own lagged values and current and lagged explanatory variables. The models were estimated by the classical method of least squares.[9]

1.3. Empirical verification of thesis 1

The monthly profitability of loans by KNF is determined by the monthly average WIBOR, as most loans are indexed to WIBOR. The time-series of variables are cointegrated; therefore the postulated relationship was determined using the ECM model.

The long-term form of the model:

$$RK_t = \beta_0 + \beta_1 WB_t + \xi_t.$$

ECM model of the form:

$$\Delta RK_t = \gamma_1 \left(RK_{t-1} - \beta_0 - \beta_1 WB_{t-1} \right) + \gamma_2 \Delta WB_t + \gamma_{11} D11_t + \gamma_{13} D13_t + \varepsilon_t,$$

where:

ΔRK_t – change in the profitability of loans by KNF (in %),

ΔWB_t – change in the monthly WIBOR (%).

The testing and estimation of the first cointegrating relationship (long-term) showed that the profitability of loans by KNF and the monthly WIBOR are cointegrated and in this case random components can be regarded as a deviation from the long-term path. This means that regardless of the passage of time, the analysed variables remain at a constant distance from each other, and the model remains just as accurate and the model error (measured by residuals) does not grow. In this case, the estimated structural parameters forming a cointegrating vector are interpreted as parameters of a long-term relationship of the form (appendix, Table 2). The interpretation of parameters of a long-term relationship indicates that if the average WIBOR is zero, the average profitability of loans by KNF would amount to 3.08% with an average estimation error of 0.15%.

[9] The calculations were performed using the Microfit 5.0.package.

If the average WIBOR increased by 1%, the average profitability of loans by KNF in the long term would increase by 0.65% with an average estimation error of 0.04%, c.p.

In the process of a quantitative verification of the model (appendix, Table 1) it was found that the variance of the random component is constant over time (BP test), the distribution of the random component is not compatible with normal distribution (JB test) and the DW test does not determine whether there is an autocorrelation of the random component of order 1. The coefficient of the determination of 78.9% indicates a fairly good adjustment of the estimated model to the actual data. The t-Student test has shown that both structural parameters are statistically significantly different from zero (appendix, Table 2). Table 3 in the appendix and Figure 1 of the appendix provide detailed statistics of the decomposition of the residual component of the model representing a deviation from the long-term path.

According to the designed study, in the second stage the ECM model was estimated using OLS (ordinary least squares). According to Granger's theorem on representation, if the variables are cointegrated they can be presented in the form of the ECM model, as there must be an adjustment mechanism, providing stationarity of the random component in long-term relationships (appendix, Table 5). The estimated value of the parameterγ_1 is negative, which indicates that the mechanism for correcting disorders is working and thus the variable average profitability of loans by KNF has returned to a state of balance after the disturbances caused by external impulses. The value of the rate of return to balance is equal to –0.81 and means that 81% of the deviation is corrected after one month; however, this trend can be regarded as durable and resulting from the fundamental characteristics of the economy. If WIBOR increases by 1 p.p., then in a short period of time, the average profitability of loans by KNF will increase by an average of 0.96 p.p. with an average estimation error of 0.28 p.p., *c.p.* In September 2011, the average profitability of loans by KNF was significantly lower than the average level by 0.53 p.p., while in December 2013 it was significantly higher by 1.69 p.p.

In the process of a quantitative verification of the model (appendix, Table 4), it was found that the variance of the random component is constant over time (BP test), the distribution of the random component is compatible with normal distribution (JB test) and the DW test determines that there is no autocorrelation of the random component order 1. The coefficient of the determination of 62% indicates a fairly good adjustment of the estimated model to the actual data. The t-Student test has shown that all structural parameters of the model are statistically significantly different from zero (appendix, Table 5).

1.4. Empirical verification of thesis 2

The monthly rate on deposits by KNF (representing the annualized quotient of net interest income and volume) is determined by the average rate on deposits by NBP (which determines the interest rates on deposits in total). The exact cost of liabilities is observed by SMRP (showing the marginal cost of the negotiated deposits).

When analysing the properties of these time series, it was found that they are non-stationary (integrated of the same order 1), but at the same time it was found that they are not cointegrated, therefore, it is not possible to apply the ECM model. In the case of modelling this relationship, the dynamic autoregressive distributed lag model (ADL) will be applied in the following form:

$$\Delta RD_t = \gamma_0 + \gamma_1 \Delta RD_{t-1} + \ldots + \gamma_p \Delta RD_{t-p} + \lambda_0 \Delta SD_t + \lambda_1 \Delta SD_{t-1} + \ldots + \lambda_q \Delta SD_{t-q} + \xi_t,$$

where:

ΔRD_t – change of monthly rate on deposits by KNF (in %),

ΔWB_t – change of monthly deposit rates by NBP (%).

The number of lags in the estimated model was set at level 2 based on the Akaike information criterion. In addition, the model included a dummy variable $D11_t$ (appendix, Table 7).

The interpretation of the parameters of the estimated ADL model indicates that if the average rate of rate on deposits by KNF increased by 1 p.p. in the previous month and return to the previous state, the average rate on deposits would decrease in the current month by an average of 0.73 p.p. with an estimation error of 0.11 p.p. A similar rise and fall two months earlier would cause a decrease of 0.30 p.p. Both parameters are statistically significantly different from zero, which means that changes in the average yields on deposits by KNF in previous periods influence the level of changes in the current period (t-Student test, Table 6). For the level of significance of 0.05, the authors established a statistically significant impact of WIBOR interest rate changes, occurring, respectively, in the previous month and two months earlier, on changes in the average yields on deposits in the current period (t-Student test, Table 6). In September 2011, the change in the average rate of yields on deposits by KNF was significantly lower than the average level, by an average of 0.82 p.p.

In the process of a quantitative verification of the model (appendix, Table 6), it was found that the variance of the random component is constant over time (BP test), the distribution of the random component is not compatible with normal distribution (JB test) and the DW test shows that there is no autocorrelation of the random component order 1. The F-test indicates a significant impact of all independent variables on the dependent variable of the model. The coefficient of determination of 58% indicates a fairly good adjustment of the estimated model to the actual data. Figure 2 (in the appendix) includes the actual and estimated values of changes in the rate on deposits by KNF.

1.5. Empirical verification of thesis 3

The monthly yields on deposits by KNF are determined by the monthly average WIBOR.

When analysing the properties of these time series, it was found that they are non-stationary (integrated of the same order 1), but at the same time it was found that they are not cointegrated, therefore, it is not possible to apply the ECM model. In the case of modelling this relationship, the dynamic autoregressive distributed lag model (ADL) will be applied in the following form:

$$\Delta RD_t = \gamma_0 + \gamma_1 \Delta RD_{t-1} + \ldots + \gamma_p \Delta RD_{t-p} + \lambda_0 \Delta WB_t + \lambda_1 \Delta WB_{t-1} + \ldots + \lambda_q \Delta WB_{t-q} + \xi_t,$$

where:

ΔRD_t – monthly yields on deposits by KNF (in %),

ΔWB_t – monthly WIBOR (%).

The number of lags in the estimated model was set at level 2 based on the Akaike information criterion. In addition, the model included a dummy variable $D11_t$ (appendix, Table 7). The interpretation of the parameters of the estimated ADL model indicates that if the average rate of cost of deposits by KNF increased by 1 p.p. in the previous month and return to the previous state, the average rate on deposits would decrease in the current month by an average of 0.69 p.p. with an estimation error of 0.10 p.p. A similar rise and fall two months earlier would cause a decrease of 0.19 p.p. Both parameters are statistically significantly different from zero, which means that changes in the average rate on deposits by KNF in previous periods influence the level of changes in the current period (t-Student test, Table 6). For the level of significance of 0.05, the authors established a statistically significant impact of WIBOR interest rate changes on changes in the average yields on deposits in the current period (t-Student test, Table 6). In September 2011, the change in the average yields on deposits by KNF was significantly lower than the average level, by an average of 0.81 p.p.

In the process of a quantitative verification of the model (appendix, Table 6), it was found that the variance of the random component is constant over time (BP test), the distribution of the random component is compatible with normal distribution (JB test) and the DW test indicates that there is no autocorrelation of the random component order 1. The F-test indicates a significant impact of all independent variables on the dependent variable of the model. The coefficient of the determination of 57% indicates a fairly good adjustment of the estimated model to the actual data. Figure 3 (in the appendix) includes the actual and estimated values of changes in the rate on deposits by KNF.

1.6. Empirical verification of thesis 4

The average monthly SMRP rate (the cost of negotiated deposits received by banks on the basis of concluded transactions) depends on the average rate of deposits by NBP.

When analysing the properties of these time series, it was found that they are non-stationary (integrated of the same order 1), and subsequently it was found that they are not cointegrated, therefore, it is not possible to apply the ECM model. In the case of modelling this relationship, the dynamic autoregressive distributed lag model (ADL) will be applied in the following form:

$$\Delta SS_t = \gamma_0 + \gamma_1 \Delta SS_{t-1} + \ldots + \gamma_p \Delta SS_{t-p} + \lambda_0 \Delta SD_t + \lambda_1 \Delta SD_{t-1} + \ldots + \lambda_q \Delta SD_{t-q} + \xi_t ,$$

where:

ΔSS_t – the average monthly SMRP rate (in %),

ΔSD_t – the average rate on deposits by NBP (%).

In the estimation of models with different combinations of lags, it has been established that none of the changes in previous periods had a statistically significant impact on changes in the average SMRP rate (appendix, Table 7). If the rate of NBP deposits increases by 1 p.p., the SMRP rate will increase by an average of 0.87 p.p. with an estimation error of 0.09 p.p. In the process of a quantitative verification of the model (appendix, Table 6), it was found that the variance of the random component is constant over time (BP test), the distribution of the random component is compatible with normal distribution (JB test) and the DW test indicates that there is no autocorrelation of the random component order 1. The F-test indicates a significant impact of the independent variable of the model on the dependent variable. The coefficient of determination of 76% indicates a fairly good adjustment of the estimated model to the actual data. Figure 4 (in the appendix) includes the actual and estimated values of changes in the average SMRP rate.

1.7. Empirical verification of thesis 5

The interest margin of the banking sector is negatively correlated with the SMRP spread due to the materialization of the basis risk. The time-series of variables are cointegrated, therefore, the postulated relationship was determined using the ECM model.

The long-term form of the model:

$$M_t = \beta_0 + \beta_1 SP_t + \xi_t .$$

ECM model of the form:

$$\Delta M_t = \gamma_1 \left(M_{t-1} - \beta_0 - \beta_1 SP_{t-1} \right) + \gamma_2 \Delta SP_t + \gamma_{13} D13_t + \varepsilon_t .$$

The testing and estimation of the cointegrating relationship (long-term) have shown that the time series of the variable margin by KNF and the SMRP spread are cointegrated and that also in the case of this model the random components can be regarded as a deviation from the long-term path (appendix, Table 2). The interpretation of parameters of the long-term relationship indicates that if the SMRP spread equalled zero, the margin by KNF would amount to 3.21% with an average estimation error of 0.14%. If the SMRP spread increased by 1%, the margin by KNF in the long term would increase by 1.65% with an average estimation error of 0.60%, *c.p.*

In the process of a quantitative verification of the model (appendix, Table 1), it was found that the variance of the random component is constant over time (BP test), the distribution of the random component is not compatible with normal distribution (JB test) and the DW test shows that there is no autocorrelation of the random component order 1. The coefficient of the determination of 20% indicates a fairly low adjustment of the estimated model to the actual data. The t-Student test has shown that both structural parameters are statistically significantly different from zero (appendix, Table 2). Table 5 in the appendix and Figure 5 of the appendix provide detailed statistics of the decomposition of the residual component of the model representing a deviation from the long-term path.

In the second stage the ECM model was estimated using OLS (ordinary least squares). The estimated value of the parameter γ_1 is negative, which indicates that the mechanism for correcting disorders is working and thus the variable margin by KNF has returned to a state of balance after the disturbances caused by external impulses. The value of the rate of return to balance is equal to –0.80 and means that 80% of the deviation is corrected after one month; however, this trend can be regarded as durable. If the SMRP spread increases by 1 p.p., then in a short period, the margin by KNF will increase by an average of 0.84 p.p. with an average estimation error of 0.44 p.p., *c.p.* In December 2013, the margin level by KNF was significantly higher than the average level by 1.23 p.p., with an estimation error of 0.16 p.p., *c.p.*

In the process of a quantitative verification of the model (appendix, Table 4), it was found that the variance of the random component is constant over time (BP test), the distribution of the random component is compatible with normal distribution (JB test) and the DW test indicates that there is no autocorrelation of the random component order 1. The coefficient of the determination of 83% indicates the high adjustment of the estimated model to the actual data. The t-Student test has shown that all structural parameters of the model are statistically significantly different from zero (appendix, Table 5).

1.8. Empirical verification of thesis 6

The average monthly SMRP rate (the cost of negotiated deposits received by banks on the basis of concluded transactions) depends on the average yields on deposits by KNF. The time-series of variables are cointegrated, therefore, the postulated relationship was determined using the ECM model.

The long-term form of the model:

$$SS_t = \beta_0 + \beta_1 RD_t + \xi_t .$$

ECM model of the form:

$$\Delta SS_t = \gamma_1 \left(SS_{t-1} - \beta_0 - \beta_1 RD_{t-1} \right) + \gamma_2 \Delta RD_t + \gamma_3 \Delta RD_{t-1} + \gamma_4 \Delta SS_{t-2} + \varepsilon_t ,$$

where:

ΔM_t – change of the monthly average SMRPrate (in p.p.),

ΔRD_t – change of the monthly rate on deposits by KNF (in p.p.).

The testing and estimation of the cointegrating relationship (long-term) has shown that the time-series of the average SMRP rate and the yields on deposits by KNF are cointegrated and also in the case of this model the random components can be regarded as a deviation from the long-term path (appendix, Table 7). The interpretation of parameters of the long-term relationship indicates that if the average yields on deposits by KNF were zero, the average SMRP rate would amount to 0.39% with an average estimation error of 0.12%. If the average yields on deposits by KNF increased by 1%, the average SMRP rate in the long term would increase by 1.31% with an average estimation error of 0.08%, *c.p.*

In the process of a quantitative verification of the model (appendix, Table 1), it was found that the variance of the random component is not constant over time (BP test) and, therefore, the model was re-estimated using a generalized OLS with heteroskedasticity correction. In the re-estimated model, the distribution of the random component is not compatible with normal distribution (JB test), and the DW test shows that the model includes a positive autocorrelation of the random component order 1. The coefficient of the determination of 89% indicates a fairly good adjustment of the estimated model to the actual data. The t-Student test has shown that both structural parameters are statistically significantly different from zero (appendix, Table 2). Table 5 in the appendix and Figure 6 in the appendix provide detailed statistics of the decomposition of the residual component of the model representing a deviation from the long-term path.

In the second stage, the ECM model was estimated using OLS. The estimated value of the parameterγ_1 is negative, which indicates that the mechanism for correcting disorders is working and thus the variable average SMRP rate has returned to a state of balance after the disturbances caused by external impulses. The value of the rate of return to balance is fairly low and equal to –0.34 and means that 34% of the deviation is corrected after one month; however, this trend can be regarded as durable. If the change in the rate on deposits increases by 1 p.p. then in the current month the average SMRP rate will increase by 0.63 p.p., whereas if a single change in the yields on deposits took place in the previous period, then in the current month the average SMRP rate will increase by 0.29 p.p., respectively, c.p. The change in the level of the average SMRP rate by 1 p.p. causes its increase by 0.54 p.p. after two months, *c.p.*

In the process of a quantitative verification of the model (appendix, Table 6) with the use of OLS, it was found that the variance of the random component is constant over time (BP test), the distribution of the random component is compatible with normal distribution (JB test) and the DW test indicates that there is no autocorrelation of the random component order 1. The coefficient of determination of 57% indicates a fairly good adjustment of the estimated model to the actual data. The t-Student test has shown that all structural parameters of the model are statistically significantly different from zero (appendix, Table 5).

Conclusions

As rightly pointed out by Cabral (2013), the deterioration of net interest income is an effective indicator of growing imbalances in the banking sector, which – in the presence of increasing disparities between off-balance sheet and balance sheet exposure – can lead to banking crises. Therefore, the analysis of the sources of fluctuations in net interest income leads to the identification of risks and – in the long run – is capable of neutralizing the risks associated with the balance sheet of the banking sector.

The study presented in this article has shown the effect of the basis risk on net interest income of the banking sector. The basis risk is materialized in the absence of a correlation between reference rates used for the indexation of assets and the real cost of financing represented by the interest on liabilities. The method for verifying the thesis on the existence of the basis risk is the analysis of the cointegration of the time series relating to various interest rates quoted on the financial market.

The analysis has shown the presence of cointegration between the following series:

– return on assets by KNF data and WIBOR,
– cost of liabilities by KNF data and SMRP,
– SMRP vs. WIBOR spread and the interest margin by KNF data.

Whereas, no cointegration has been shown for:

– cost of liabilities by KNF and WIBOR,
– cost of liabilities by KNF data and the interest on deposits by NBP data,
– interest on deposits by NBP data and by SMRP data.

On the basis of the above dependencies, error correction models (ECM) were developed in the case of cointegrating dependence, and autoregressive distributed lag models (ADL) were developed in the absence of cointegration. The determined statistical properties determination and the econometric models constructed allow formulating the following conclusions:

1. The return on assets of the banking sector is naturally determined by the level of WIBOR, as in the case of most long-term contracts, WIBOR is a key component of the calculation of cash flows from floating rate payments.
2. The cost of liabilities is not determined by WIBOR, as the cost of financing differs from WIBOR rates and is not correlated with them. The cost of liabilities is directly affected by interest rates on the deposits collected by banks, with particular emphasis on sustainable forms of financing i.e. deposits of individual and corporate customers.
3. Despite the lack of cointegration between interest rates on negotiated term deposits and total deposits (which results from the different price dynamics of these two types of financing, due to the high proportion of short-term deposits and non-negotiated deposits (table-based) in the financing of banks), the authors demonstrated the long-term impact of interest rates on negotiated deposits (by SMRP) on the cost of the banks' liabilities.

4. The authors proved a significant long-term dependence determined by the difference between the average monthly SMRP rate and WIBOR (jointly for the maturities from 1M to 6M) on net interest income of the banking sector was determined.

The last conclusion is crucial. It indicates that the observed-to-date prices of negotiated deposits are a particular kind of indicator of the current marginal cost of financing of the banking sector, which translates directly into the volatility of the cost of bank liabilities. On the other hand, the reference rate of the interbank market (WIBOR), "attached" to loan agreements, determines, to a large extent, the profitability of the banking sector assets. The divergent dynamics of SMRP and WIBOR rates represents the materialization of the basis risk arising from various changes in yields on assets and liabilities. This situation causes fluctuations in net interest income. In an environment of falling interest rates (which is largely covered by the sample studied), the basis risk means a faster decline in interest rates on loans than the cost of financing the loan portfolio, which means a deterioration of net interest income of the sector. The higher the level of SMRP price deviations from WIBOR, the greater the risk of instability of the difference between the interest income and cost, which results in the lower profitability of the banking sector.

The divergent dynamics of the SMRP rate and WIBOR results from the different nature of these interest rates. WIBOR, to a large extent, expresses the expectations as to the changes of interest rates in the future, as it constitutes the underlying instrument for FRA and IRS[10] derivatives contracts, where the turnover is much higher than on the interbank deposit market. Since the financial crisis in 2007–2009, due to the disappearance of the market of unsecured interbank deposits for periods longer than 1 week, the reference rate for the periods used in the loans market and the derivatives market has not had a significant transactional database. Thus, its sensitivity to the liquidity situation and the creditworthiness of the banking sector is minimal. At the same time, due to the increasing role of the CRD IV regulation and the lack of credit lines for interbank deposits, the wholesale market, which is the basis for determining the index, is becoming less significant. In fact, financing from banks is inefficient from the regulatory point of view – according to the standards defined in regulations; deposits from financial institutions cannot be regarded as a stable source of funding.

In this situation, the non-bank deposit market is becoming the main source of bank funding. The measure of the marginal price on this market is transactional data recorded in SMRP, which take into account the current liquidity situation, creditworthiness and competition in the banking sector. Thus, the dynamics of reference rates differs from the volatility of deposit

[10] It is worth noting that in the case of the major currencies, an IBOR-type rate is no longer used to determine the factors discounting future cash flows. Even the instruments, for which cash flows are determined by IBOR, use OIS rates to calculate the net present value, due to the widespread collateralization of off-balance sheet items.

prices, which, in turn, show inertia resulting from strong competition for customer funds under high Bank Guarantee Fund thresholds neutralizing the effect of varying degrees of reliability of particular banks.

Appendix

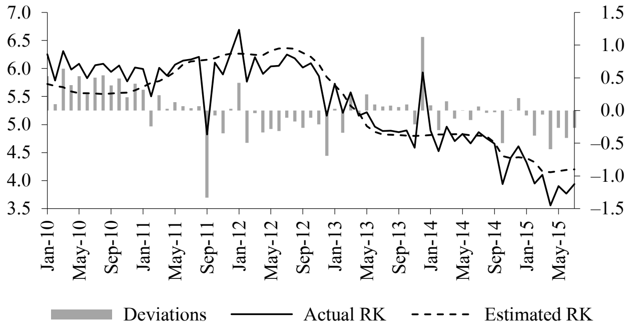

Figure 1. Actual and estimated profitability of loans by KNF and the deviations from the long-term path

Source: own calculations.

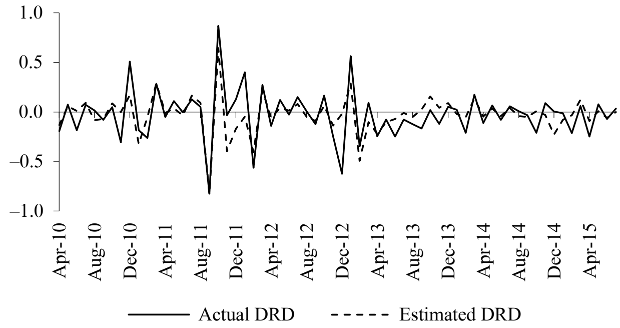

Figure 2. Actual and estimated change in the rate on deposits by KNF

Source: own calculations.

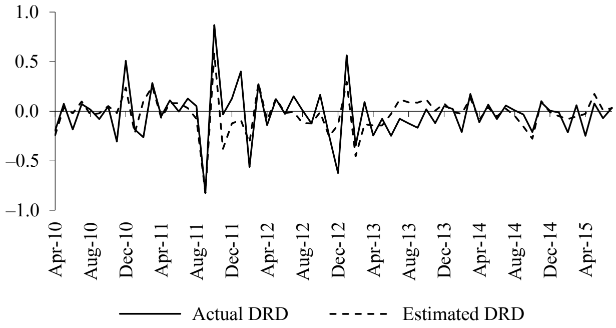

Figure 3. Actual and estimated change in the rate on deposits by KNF

Source: own calculations.

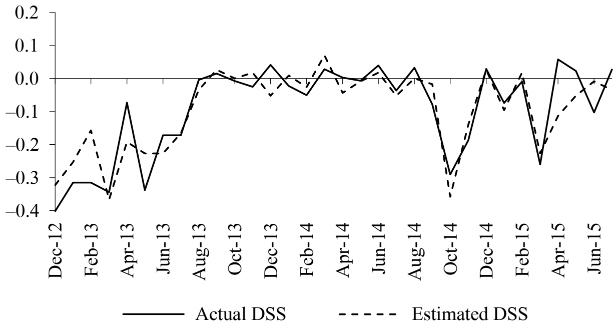

Figure 4. Actual and estimated change in the average SMRP rate

Source: own calculations

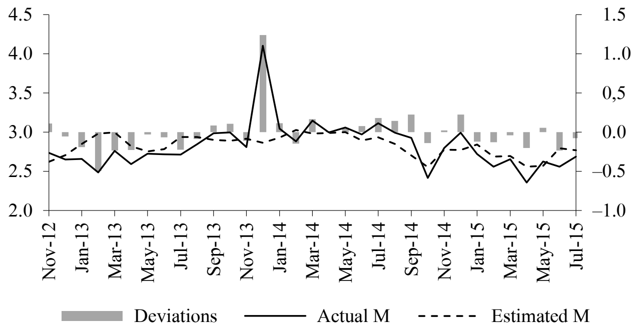

Figure 5. Actual and estimated margin by KNF and deviations from the long-term path

Source: own calculations.

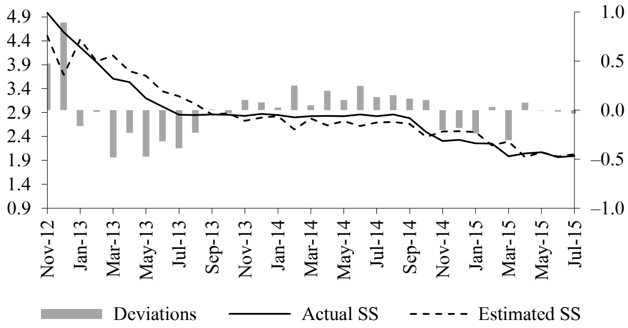

Figure 6. Actual and estimated average SMRP rate and deviations from the long-term path

Source: own calculations.

Table 1. Diagnostic tests of OLS estimates of long-term relationships

Model	T	DW	JB	BP	R^2
1	67	1.46*	16.29*	0.85	0.79
5	33	1.73	197.35*	1.48	0.20
6	33	1.00*	10.69*	-	0.89

* The null hypothesis was rejected for the level of significance of 0.05.

Source: own calculations.

Table 2. The estimation of structural parameters of long-term relationships

Model	β_0	SE(β0)	β1	SE(β1)
1	3.08*	0.15	0.65*	0.04
5	3.21*	0.14	−1.65*	0.60
6	0.39*	0.12	1.31*	0.08

* The null hypothesis was rejected for the level of significance of 0.05.

Source: own calculations.

Table 3. Descriptive statistics of long-term relationship deviations

Statistics	Model 1	Model 5	Model 6
Average	0.000	0.000	–0.008
Median	0.025	0.039	0.005
Minimum	–1.333	–0.496	–0.482
Maximum	1.126	1.238	0.892
Standard deviation	0.363	0.275	0.271
Skewness coefficient	–0.260	2.657	0.854
Kurtosis	2.359	10.737	2.203

Source: own calculations.

Table 4. Diagnostic tests of OLS estimations of ECM models

Model	T	DW	JB	BP	R^2
1	66	1.84	1.98	10.57	0.62
5	32	1.77	0.49	1.09	0.82
6	30	1.81	0.55	5.88	0.57

The Durbin-Watson (DW) statistics was used to study the autocorrelation of the random component of the model of order 1, the Jarque-Bera (JB) statistics was used to study the normality of distribution of the random component of the model, the Breusch-Pagan (BP) statistics was used to study the stability of the variance of the random component of the model, R^2 coefficient of determination.

Source: own calculations.

Table 5. The estimation of structural parameters of EMC models

Model	γ_1	SE(γ_1)	γ_2	SE(γ_2)	γ_3	SE(γ_3)	γ_4	SE(γ_4)	γ_{11}	SE(γ_{11})	γ_{13}	SE(γ_{13})
1	−0.81*	0.10	0.96*	0.28	–	–	–	–	−0.53*	0.13	1.69*	0.31
5	−0.80*	0.10	−0.84**	0.44	–	–	–	–	–	–	1.23*	0.16
6	–0.34*	0.14	0.63*	0.21	0.29*	0.13	0.54*	0.14				

* The null hypothesis was rejected for the level of significance of 0.05.

** The null hypothesis was rejected for the level of significance of 0.10.

Source: own calculations.

Table 6. Diagnostic tests of estimations of ADL models

Test	Model 2	Model 3	Model 4
T	64	64	32
h	1.48	1.25	–
DW	–	–	2.30
JB	14.03*	2.77	0.23
BP	7.55	2.50	2.11
F	15.79*	24.74*	95.05*
R^2	0.58	0.57	0.76
t(γ_0)	–0.83	–0.35	–1.18
t(γ_1)	–6.89*	–6.64*	–
t(γ_2)	–2.78*	–1.90**	–
t(λ_0)	–	4.16*	9.75*
t(λ1)	2.37*	–	–
t(λ_2)	2.05*	–	–
t(λ_{d11})	–4.64*	–4.40*	–

The h-Durbin and Durbin-Watson (DW) statistics were used to study the autocorrelation of the random component of the model of order 1, the Jarque-Bera (JB) statistics was used to study the normality of distribution of the random component of the model, the Breusch-Pagan (BP) statistics was used to study the stability in time of the variance of the random component of the model, the F-test – to study the total significance of the structural parameters of the model, R^2 coefficient of determination, t (*) – t test statistics for testing the significance of structural parameters of the model.

* The null hypothesis was rejected for the level of significance of 0.05.

** The null hypothesis was rejected for the level of significance of 0.10.

Source: own calculations.

Table 7. The estimation of structural parameters of ADL models

Model	γ_0	SE(γ_0)	γ_1	SE(γ_1)	γ_2	SE(γ_2)	λ_0	SE(λ_0)	λ_1	SE(λ_1)	λ_2	SE(λ_2)	γ_{11}	SE(γ_{11})
2	–0.02	0.02	–0.30*	0.11	–	–	–	–	0.44*	0.19	0.37*	0.18	–0.82*	0.18
3	0.01	0.02	–0.69*	0.10	–0.19*	0.10	0.76*	0.18	–	–	–	–	–0.81*	0.18
4	–0.02*	0.01	–	–	–	–	0.87*	0.09	–	–	–	–	–	–

* The null hypothesis was rejected for the level of significance of 0.05.
** The null hypothesis was rejected for the level of significance of 0.10.

Source: own calculations.

References

Brousseau, V., Chailloux, A., Durré, A. (2013). *Fixing the Fixings: What Road to a More Representative Money Market Benchmark?* IMF Working Paper No. 13/131, May 29.

Cabral, R. (2013). A perspective on the symptoms and causes of the financial crisis. *Journal of Banking & Finance*, *37* (1), 103–117.

Charemza, W.W., Deadman, D.F. (1997). *Nowa Ekonometria*. Warszawa: PWE.

Demirgüç-Kunt, A., Huizinga, H. (1999). Determinants of commercial bank interest margins and profitability: some international evidence. *The World Bank Economic Review*, *13* (2), 379–408.

Engel, R.F., Granger, C.W.J. (eds.) (1991). Long Run Economic Relation, Readings in Cointegrations. Oxford: Oxford University Press.

English, W.B. (2002). Interest rate risk and bank net interest margins. *BIS Quarterly Review*, *12* (2), 67–82.

Ho, T.S., Saunders, A. (1981). The determinants of bank interest margins: theory and empirical evidence. *Journal of Financial and Quantitative Analysis*, *16* (4), 581–600.

Maes, K. (2004). Interest rate risk in the Belgian banking sector. *Financial Stability Review*, *2* (1), 157–179.

Ötker-Robe, İ., Pazarbasioglu, C. et al. (2010). *Impact of regulatory reforms on large and complex financial institutions*. IMF Staff Position Note, November 23,

Peng, W. et al. (2003). The impact of interest rate shocks on the performance of the banking sector. *HKMA Quarterly Bulletin*, *35*.

Welfe, A. (ed.) (2013). Analiza kointegracyjna w makromodelowaniu. Warszawa: PWE.

www.nbp.pl.

www.knf.gov.pl.

www.smrp.pl.