

NBP Working Paper No. 227

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Michał Markun – Narodowy Bank Polski, Domestic Operations Department, and Cardinal
Stefan Wyszyński University in Warsaw; michal.markun@nbp.pl
Anna Mospan – Narodowy Bank Polski, Domestic Operations Department, and a Ph.D.
student at Łódź University of Technology; anna.mospan@nbp.pl

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Abstract

The present paper examines the term premia in the interbank money market in Poland. We use analyst surveys to proxy interest rate expectations and forward rate agreement (FRA) market data to construct term premia. We consider the term premia at shorter and longer horizons. Both premia follow autoregressive, stationary processes of low orders. The longer term premium is higher and more volatile than the shorter one; moreover, it is also characterized by substantially higher persistence. Our findings provide direct evidence against the efficient markets hypothesis (EMH) at the short end of the Polish yield curve and indicate areas of potential ineffectiveness of the monetary policy transmission mechanism.

Keywords: short-term interest rate, expectations, term premium, persistence, surveys, Poland

JEL Classification: C83, E43, E58, G23

1. Introduction

Term premia in financial markets are important from many points of view. In academic research time-varying term premia, revealed in their persistence, are invoked to explain the rejections of the EMH. For central banks, their characteristics may help in explaining the shape and the evolution of the term structure of interest rates, thus revealing information useful in decision-making. Understanding the issue of term premium is also crucial because its persistence may potentially interfere with the efficient transmission of monetary policy impulses to the real economy. Indeed, one instance of such a phenomenon, known as the ‘Greenspan conundrum’ – the fact that the 10-year Treasury yields failed to increase despite a 150 basis point hike in the federal funds rate in 2005, was one of the factors that drew researchers’ attention to developing modelling frameworks and estimation methods of term premia and contributed to the growth of the relevant literature. While the ‘Greenspan conundrum’ refers to long-term interest rates and to the term premium itself rather than its persistence, the issue of the term premium persistence is probably no less relevant for short-term yields, which directly affect economic agents and over which monetary policy authorities believe to have immediate control.

There are several interrelated definitions of term premia. The most convenient in our context is the one based on forward rates inherent in the yield curve, which can be conceptually decomposed into two parts: an expectational term and a term premium. The term premium, thus, is by definition nothing but a difference between the forward yield and the corresponding expected interest rate. Therefore, if one wishes to investigate the term premium, the critical step is the construction of the expected rate. One way, often used in empirical research, is to rely on economic theory by applying the rational expectations principle. Alternatively, if available, the expected rate can be read directly from surveys among financial markets participants. There are several important advantages of the surveys that make for their growing use in the literature: (i) they provide an observable proxy for market expectations about future rates in real-time, (ii) it is a non-estimated and model-free variable, thus saving us from additional estimation and model uncertainty, (iii) surveys are robust to learning and easily accommodate structural breaks in the data. Howev-

er, the use of survey measures may also be criticized on several grounds: (i) they include noise and their use hinges on the identifying assumption that they measure market expectations correctly, (ii) it may be unclear which measure of central tendency is provided by survey participants, (iii) survey participants make their predictions at different moments in time, which are thus based on different information sets, and may give strategic forecasts rather than their true expectations, (iv) lower data frequency than financial markets' and usually only short horizons covered, (v) several specific problems in the context of testing for bias in survey forecasts as a test for the EMH, which derive from the way survey data are aggregated, from the use of particular data releases for testing, and from the dependence of the forecast on the individual utility function, among others.

As will be explained later, any persistence measure is a summary statistic of a certain infinite-dimensional vector introduced to facilitate its interpretation. It captures the idea that a process responds gradually to shocks or that it remains close to its recent history. It thus involves a degree of abstraction, which implies that using various measures is needed to have a good understanding of the phenomenon. We draw mainly on a growing literature on inflation persistence in the context of a standard univariate time-series representations, namely autoregressions, which we apply to the term premia. The measures we employ are mostly simple functions of the model's parameters. Additionally, we include two non-parametric measures, and the so-called half-life, which has a particular presentational appeal since the unit in which it is expressed is time periods.

The objections against surveys notwithstanding, in this paper we investigate to what extent the information contained in the surveys can shed light on the issue of term premia persistence in the Polish interbank money market. We are not aware of any previous paper investigating term premium persistence in Poland, and thus of any paper to compare to. Since there is a lack of empirical studies concerning the term premium of emerging markets, all the more so in the case of Poland, our analysis is a novel contribution to the literature. To our knowledge, this paper constitutes also the first attempt at an analysis of the term premium constructed from market data and survey expectations for Poland. We use Thomson Reuters polls on the

WIBOR 3M rate as well as quotations on FRA contracts as market predictors of the WIBOR 3M rate on the Polish market; the FRA market is a very liquid segment of the Polish financial market, thus efficiently aggregating investors' differing outlooks on the short-term interest rate. As we will explain later we are able to investigate two series of term premia, one at a shorter horizon and another at a longer horizon. We acknowledge that building a system for forward premia across a spectrum of time horizons would be an interesting problem on its own, but due to data limitations and also because this would take us too far away from the aim of this paper, we do not model the term structure of forward premia. Since we are interested in the modeling of the conditional expectation of the data generating process of the term premia, as opposed to, e.g. conditional variance, we first implement a detailed cointegration analysis of the system consisting of the above two variables from which the premia are derived. We then present various persistence measures used in the literature to investigate persistence of a wide variety of economic variables. The present paper is the first case of applying them to term premia. The results of our analysis show that the longer-horizon term premium is not only higher and more volatile than the shorter-horizon term premium, but also more persistent, as corroborated by all the persistence measures. Both are, however, stationary processes – which is to be expected, after all they are both supposed to express the underlying and unobserved expectations – and moderately correlated with each other.

The remainder of the paper is structured as follows: Section 2 presents an overview of the studies related to the use of surveys as a proxy for market rate expectations and the Polish interbank market as well as works concerning the issue of persistence in time-series; Section 3 explains the economics of a FRA contract and the development of the FRA market in Poland; Section 4 describes the applied methodology; Section 5 presents the details of data construction and discusses empirical results concerning term premia persistence; the final section contains the conclusions.

2. Literature overview

Since the fundamental works by Samuelson (1965) and Fama (1970), the EMH has been studied intensely, but despite huge theoretical literature and numerous empirical studies, the dispute over the validity of the hypothesis remains unresolved. We refer the reader to a review paper by Lo (2007) and two works by Shiller (2003) and by Malkiel (2003), in which authors made a comprehensive analysis of the available results, presented theoretical models and empirical methods of testing the EMH, and finally summarized criticism and the current state of the debate, concluding that the consensus among economists had still not been reached.

The expected rates, which are crucial in the research on the EMH, may either be constructed or inquired about among market participants. The use of surveys to proxy market expectations of interest rates was pioneered in the works by Kane and Malkiel in the sixties (see e.g. Kane and Malkiel, 1967).¹ In our context it is important to mention Friedman (1980), who showed the bias of subjective predictions of market participants and the lack of informational efficiency in the case of long-term interest rates. Survey-based tests were used in numerous other papers more generally as an evidence for (or against) rationality of expectations.² Friedman's results were refuted by Mishkin (1981), who questioned the usefulness of surveys as a reflection of market participants' expectations. Webb (1987) provided a more comprehensive analysis of the issue, concluding that the use of surveys may result in inaccurate information about efficiency or rationality of expectations. In a model with the spread between long and short rates as a predictor, Froot (1989) used survey expectations and decomposed the prediction bias for interest rates into a part attributable to expectational errors and a component connected to term premia; he thus overcame the joint hypothesis (i.e. efficient markets and rational expectations) problem besetting previous tests and concluded that at short maturities it is the EMH that fails³. While the early works that used surveys were concerned mainly with testing the EMH and term premia were not analysed *per se*, their economic importance

¹ The earliest reference we were able to find goes back to Wallich (1946).

² In the literature there are also other properties of rational expectations than just unbiasedness or informational efficiency, referring, e.g. to convergence or volatility.

³ The bias of the long-maturity contracts was caused by systematic expectation errors.

led over time to a rise in research on their dynamics and determinants. As a recent example, Dick et al. (2013) resorted to surveys in the construction of expected term premium changes and then studied their economic drivers for individual survey participants, showing by the way, to some surprise, that an aggregate measure of term premium expectations has predictive power for actual bond excess returns.

Although there is plenty of empirical literature that studies yield components using surveys for developed markets, similar investigations for emerging markets are scarce. One of those rare examples is the work by Horváth et al. (2014), who examined the term premium calculated with survey predictions and market data for fifteen emerging countries jointly (using principal components analysis to extract the common factor), including Poland. Their work paid special attention to Hungarian yield components. Their results indicated a significant reaction of the emerging market term premia to main global news (e.g. ECB or Fed communications). However, this dependency was not observed in the case of the Hungarian term premia, which reacted more to domestic events.

Moving to Poland, a comprehensive overview on the EMH and rational expectations can be found in Tomczyk (2004, 2011). Other works include Kluza and Sławiński (2003), who observed conditions for arbitrage between the bond and the FRA markets and Włodarczyk (2008), who analysed the FRA market reaction to the Polish Monetary Policy Council (MPC) communication and observed that the results were different for different parts of the yield curve and time horizons of the contracts. The reaction of the particular financial instruments was an argument against the EMH. The term structure of the Polish interbank market was also investigated by Konstantinou (2005), Bruzda et al. (2006), Blangiewicz and Miłobędzki (2009) and Kliber and Pluciennik (2011). Konstantinou (2005) studied the short end of the yield curve and one of his conclusions was that the actual yield spread indeed incorporated important information for changes of the future interest rates and could predict changes of the interest rates. Blangiewicz and Miłobędzki (2009) found much support for the rational expectation hypothesis, which involves informational efficiency, with the time-varying term premium. Bruzda et al. (2006) came to the conclusion that policy-makers had no influence on the performance of the yield spreads of the

Polish term structure. Kliber and Płuciennik (2011) examined the Polish interbank market risk premium, which turned out to be small and time-varying, though its volatility reacted significantly to the MPC meetings. None of the above papers on Poland used survey data or term premia constructed upon them.

The persistence of term premia, to be formally defined in Section 4, is of key importance, as explained in the Introduction. Intuitively, a series is considered to be persistent if it shows a tendency to stay near where it has been recently, provided there are no forces that move it away. Batini and Nelson (2002) formulated three measures of inflation persistence. In addition, many other different measures, to be presented further in our article, were used to analyse persistence in various contexts. Cochrane (1988) used the size of a random walk in a series to measure persistence of the U.S. gross national product, which turned out to have little long-term persistence. It was then extended to the capital markets by Vošvrda (2006), who showed high level of persistence of several equity indices, thus concluding a failure of the EMH for these capital markets. In the context of interest rates, persistence and mean-reversion were analysed, e.g. in Lai (1997). Robalo Marques (2004) and Dias and Robalo Marques (2005) described various existing measures of persistence, proposed yet another of their own, applied them to inflation and pointed out that the degree of inflation persistence depends on the type assumption made about the mean (constant or time varying) of the process. The half-life became widespread in the economic literature – originally in dealing with the issue of the purchasing power parity. References include the classic paper by Rogoff (1996) and more recently Murray and Papell (2005), Caporale et al. (2005), and Rossi (2005). However, the concept was also employed to inquire about the persistence of commodity markets, see Cashin et al. (1999).

3. Forward Rate Agreement

FRA definition

A FRA is a contract where the two parties (the buyer and the seller of the contract) agree on a future exchange of cash flow based on a fixed rate (defined in the contract) and a floating rate – the so-called reference rate (IBOR-type rate). The buyer of the FRA is required to pay to the seller on a settlement day if the reference rate (from a fixing day) is lower than the contract rate; otherwise the seller of the contract pays to the buyer. The paid amount is discounted to the settlement date.

The general illustration of the entire life of the FRA $Y \times Z$ is shown in Figure 1. Y presents an initial period lasting from the contract conclusion till the settlement date and Z is a period from the contract conclusion till the maturity day, an expression $(Y - Z)$ defines the lead time of the FRA.

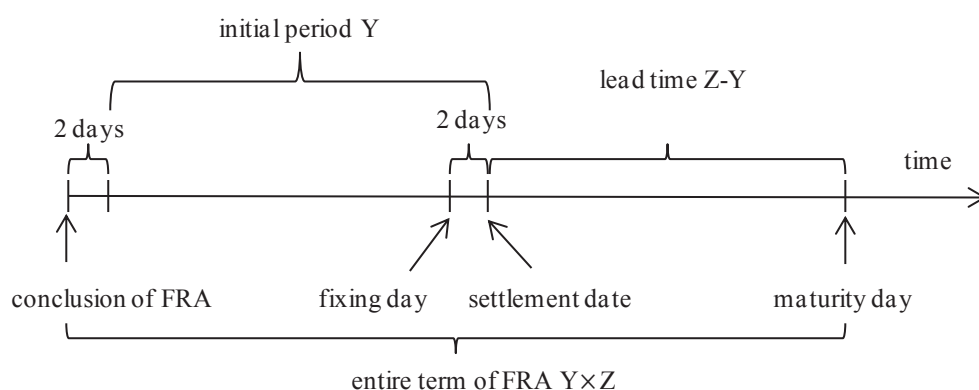


Figure 1. The entire life of a FRA $Y \times Z$ contract.

The paid amount is determined by the following formula:

$$Paid\ amount = \frac{(r_{float} - r_{fixed}) \cdot d / N}{1 + r_{float} \cdot d / N} V, \quad (1)$$

where: r_{fixed} – fixed interest rate agreed by the contract, r_{float} – settlement interest rate (IBOR), V – notional principal of the contract, d – term to maturity of the contract in days, N – number of days in the base year used in the interbank money market for deposits in the currency of the FRA contract (for the PLN FRA, 360 or 365 days are possible; on many other markets the FRAs are quoted on a $d/365$ basis). Additionally, there is no margin in the FRAs, contrary to the futures contracts.

The FRA contracts are commonly used for two purposes: (i) as a hedge against undesirable movements in short-term interest rates (i.e. they allow institutions to lock in future interbank borrowing rates), or (ii) to express IBOR-type rate expectations (assuming the EMH).

Polish FRA market

The over-the-counter (OTC) market in Poland is decentralized in nature, and its main participants are banks. At the end of 2013 there were 40 banks and branches of credit institutions that reported above PLN 1.7 trillion of gross nominal value of off-balance-sheet positions from operations in the OTC market (NBP, 2014). Approximately 87% of the gross off-balance-sheet positions were held in interest rate derivatives, mainly in IRS and FRA, 55.6% and 26.4% of all interest rate derivatives respectively. Moreover, PLN-denominated instruments prevailed and constituted about 90% of the whole market.

The FRA market denominated in PLN was formed in the second half of 1998. Initially, it involved only domestic banks, though later London's banks joined the market. Up till now, the banks have predominantly engaged in FRA transactions denominated in PLN; the share of other currencies in FRA transactions in 2013 was less than half a percent.

One-month (1M), three-month (3M) and six-month (6M) WIBOR rates are the reference rates in the Polish FRA market. The values of operations under the FRAs in the Polish interbank market in 2013 were PLN 500 and 300 million settled at 1M and 3M WIBOR rates correspondingly, and PLN 150 or 200 million settled at the 6M WIBOR rate.

The segment of PLN-denominated FRA transactions in is the most developed segment of the OTC market in Poland, with about PLN 5.8 billion of the average daily net turnover in 2013. The global financial crisis affected the development of the Polish financial system and was reflected by a decrease in derivatives turnover. Thus the average turnover of the FRAs declined by almost 70% during 2009 (see Figure 2). Even three years after the crisis, the FRA market did not come back to levels of 2007-2008. Similarly, the turnover of IRS and OIS declined by about 62% and 54% correspondingly in 2009. After the global financial crisis, the share of the

FRA transactions with maturities exceeding 9 months declined by almost half and stayed close to this level over the following years. At the same time the share of transactions with 6-9M maturities increased more than one and a half times.

Since the second quarter of 2012, the FRA market activity has been gradually increasing owing to the growth of expectations of the central bank rate cuts reflected in these instruments. The share of the five most active banks in the FRA market turnover has slightly decreased during the last thirteen years, though it remained above 80% overall. The average daily value of the FRA transactions between domestic banks in 2013 was to the tune of 1.2 billion PLN, while with non-bank institutions it was only PLN 92.4 million. The largest daily average value was recorded for transactions of domestic banks with non-residents – PLN 4.4 billion. The term structure of FRA transactions in 2013 turned out to be more balanced. The share of the short end (less than 1M) of the FRA term structure grew in 2013, while the share of the contracts with maturities longer than nine months decreased slightly (see Figure 3). The segment of the FRA market exceeding one year is currently presented by 12x15, 12x18, 15x18, 15x21, 18x21, 18x24 and 21x24 FRAs.

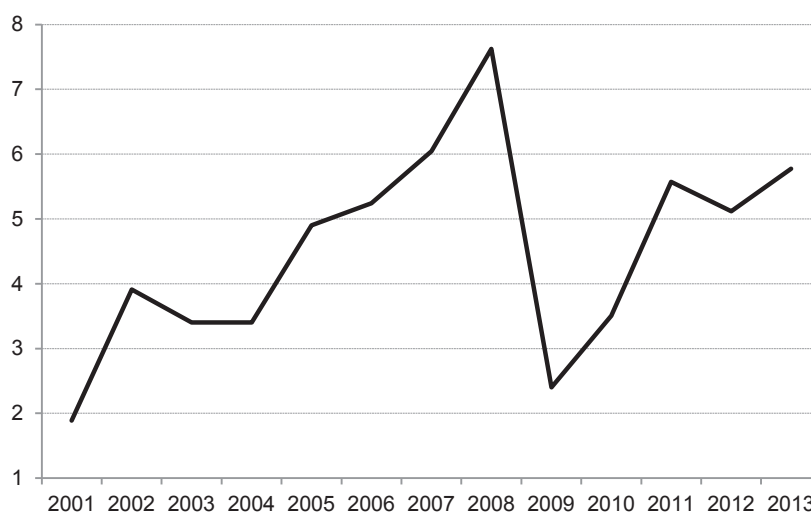


Figure 2. The average daily net turnover in FRA contracts in the Polish financial market between 2001 and 2013 (PLN billion).

Source: Own analysis based on the financial system development reports of Narodowy Bank Polski.

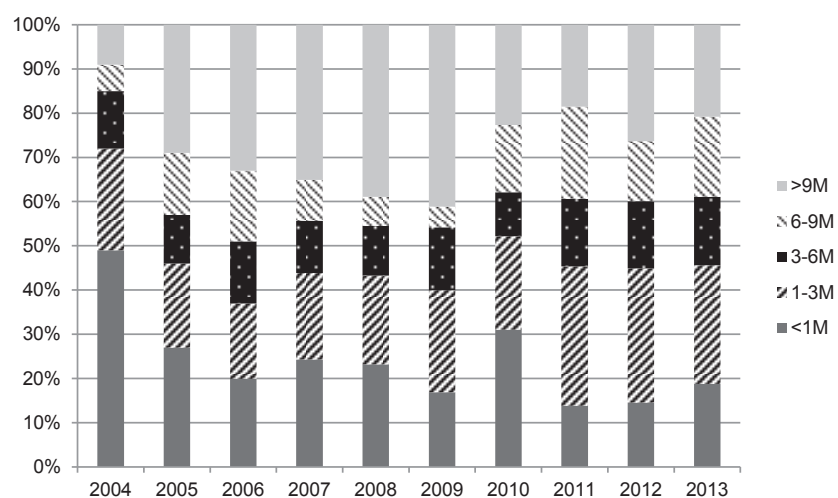


Figure 3. The term structure of PLN denominated FRA transactions between 2001 and 2013.

Source: Own analysis based on the financial system development reports of Narodowy Bank Polski.

4. Methodology

One aim of the paper is to provide a satisfactory statistical description of the data generating process of the term premium, rather than to seek explanations for its behaviour, for which we hope our initial analysis will serve as a starting point. Since there is no theory or consensus on the modelling framework appropriate for such data, we opted for standard autoregressive models methodology to be found in any textbook on time-series econometrics. It consists of (i) testing for unit roots in univariate series of FRA quotations and survey expectations, (ii) fitting a bivariate VAR system and (iii) determining cointegration rank, (iv) imposing restrictions on the cointegration space and, potentially, on the short-term model dynamics. Diagnostic checks are applied at each step to see if the chosen model fits the data. Due to data availability, we apply the above procedure to bivariate systems at a shorter and longer horizon. What we get in each case is a single stationary linear combination of short-term rate forwards and expectations, restricted to be interpreted as the respective term premium; the implied model for the premium is therefore a univariate autoregression. Alternatively we could construct term premia by referring to the definition of the term premium, directly by taking the adequate difference; however, we believe that the construction through cointegration analysis, which we consider yet another contribution of this paper, has sounder methodological foundations. Secondly, having in mind the central question of this paper, we concentrate on providing interpretation of the results in terms of term premium persistence using a number of persistence measures.

Persistence, after Robalo Marques (2004), is most often formally defined as the speed with which a stationary process converges to its long-run equilibrium, or the mean, after a shock. This will be the approach to persistence taken in the present paper, the interest lies thus in mean-reversion – if the speed of convergence to the equilibrium after a shock is low, we say that the process is persistent, otherwise it is non-persistent. As we recalled in the literature overview, Batini and Nelson (2002) proposed three definitions of persistence specifically for inflation: (i) “positive serial correlation in inflation”, (ii) “lags between systematic monetary policy actions and their (peak) effect on inflation”, and (iii) “lagged responses of inflation to non-

systematic policy actions (i.e. policy shocks)”. Measures quantifying persistence of shocks in economic time series can also be found in earlier literature (e.g. Cashin et al., 1999). Different models may lead to specific measures of persistence, but the idea of a process responding gradually to shocks or remaining close to its recent history should always be preserved.

The concept of persistence is most easily understood in the framework of univariate autoregressions. For the sake of presenting the interpretation of the concept, we shall assume that the term premium follows a stationary autoregressive process of order p , denoted AR(p), which is written as:

$$y_t = \alpha + \sum_{i=1}^p \beta_i y_{t-i} + \varepsilon_t, \quad (2)$$

where α is a constant, β_i are coefficients of the autoregression, ε_t is an unobservable white noise process with zero mean and time invariant variance. The equation can be reparametrized as:

$$\Delta y_t = \sum_{i=1}^{p-1} \delta_i \Delta y_{t-i} + (\rho - 1)[y_{t-1} - \mu] + \varepsilon_t, \quad (3)$$

where $\rho = \sum_{j=1}^p \beta_j$, $\delta_i = -\sum_{j=i+1}^p \beta_j$ and $\mu = \alpha/(1 - \rho)$. Stationarity, the assumption that we will test for, implies that $|\rho| < 1$. From the adopted modelling framework and the definition of persistence there follows the crucial role of the autoregressive coefficients for the speed of the propagation of a shock. Suppose a series has stabilised at the equilibrium, but a positive shock ε_t is realised at time t . Then the series is above its mean and the deviation $(y_{t-1} - \mu)$ will contribute as a driver to a negative change of the series in the following period. The strength of this contribution bringing the series closer to its mean depends on the coefficient $(\rho - 1)$, which in turn depends on the autoregressive coefficients β_i s. The closer ρ is to unity, the slower is the speed of mean reversion, and thus, the higher is the persistence of the process.

There are several approaches to persistence proposed in the literature. Most of them derive from the autocorrelation function of the process. Below we provide

a short listing emphasising the measures we use in the present paper, referring the reader for details to the literature reviewed above.

First, persistence is by definition closely linked to *the impulse response function* of the process, often presented as a plot, which is, however, not a useful measure as it is an infinite-dimensional vector. Similarly, *a plot of autocorrelation function* reflects the series' persistence, because the slower the pace of the decay of the autocorrelations of the process, the more persistent it is. To overcome the drawback of dimensionality, several scalar statistics have been proposed. In fact, the measures reviewed below simply provide alternative ways of quantifying, or summarising, the speed of the decay of the above plots.

Second, one popular measure of persistence is *the sum of autoregressive coefficients*, ρ , which is also monotonically related to *the cumulative impulse response* (by a formula $CIRF = 1/(1 - \rho)$) measuring persistence as the sum of the deviations from equilibrium generated during the whole convergence period, thus approximating long-run impulse response of the process to a unit shock. The choice of this parameter as a measure of persistence is obvious from equation (3) which includes $(\rho - 1)[y_{t-1} - \mu]$ on the right hand side. If at some point in time the process lies above (below) the mean μ , the positive (negative) deviation will result in a negative (positive) contribution to a change of the process in the next period, thus bringing it closer to the equilibrium. Moreover, the higher the sum, the slower the mean reversion, hence, the two concepts are closely interrelated.

From the above discussion it follows that the more often a stationary process crosses its mean, the less persistent it is. The measure that exploits this insight, proposed by Robalo Marques (2005), is *the unconditional probability of not crossing the mean in a given period* (equivalently, 1 minus the probability of mean reversion), denoted γ . The probability can be estimated non-parametrically by $\hat{\gamma} = 1 - n/T$, where n denotes the number of times the series crosses the mean during a time interval with $T + 1$ observations, and thus is expected to be immune to potential model misspecifications and robust against outliers. Estimates of γ close to 0.5 – a theoretical value for a symmetric zero mean white noise process – signal the absence of any significant persistence.

Another approach, proposed by Cochrane (1988), is to measure *the size of a random walk component in the process* and ask how large the variance of shocks to the random walk component in the series is compared to the variance of its differences. This measure, denoted by P , is operationalized by taking a limiting ratio of the k period variance to the one period variance divided by k . In short, $P = \lim_{k \rightarrow \infty} V_k$, where $V_k = \frac{1}{k} \frac{\text{var}(y_t - y_{t-k})}{\text{var}(y_t - y_{t-1})}$. P also equals the normalized spectral density function of its increments, which can be derived from the parameters of (3). Otherwise it can be estimated non-parametrically, avoiding the risk of model misspecification, by the Bartlett estimator, which is thus our method of choice. Hence, the size of the random walk is measured by:

$$P \approx 1 + 2 \sum_{i=1}^M \left(1 - \frac{i}{k+1}\right) \rho_i, \quad (4)$$

where M is the Bartlett window width and ρ_i is the autocorrelation coefficient.

Finally, although not without criticisms, *the half-life*, i.e. the number of periods required for deviations from the equilibrium in response to a unit shock to subside permanently below one half, remains one of the most popular measures in the literature on persistence. For an AR(1) process it is simply given by $h = \ln(1/2)/\ln(\rho)$, a formula which is often and incorrectly used for higher order AR processes as an approximation, which is evident in the case of non-monotonically decaying impulse responses. In the present paper, to avoid this approximation error we read h directly from the impulse response function.

There are also other measures of persistence in the literature, e.g. conventional unit-root tests, the autocorrelation coefficient of order one, the dominant root of the univariate autoregressive, the relative contributions of permanent and transitory components of the series. However, we skipped them as less frequently used and due to their inferior performance reported in Dias, Robalo Marques (2005).

5. Empirical analysis

Data

Data consist of monthly observations from February 2004 to July 2014 (126 observations in total). Market expectations of the WIBOR 3M rate were obtained as median values from Thomson Reuters⁴ polls, the benchmark for Polish financial markets. Each month the poll provides forecasts at three horizons: end-of-month, end-of-year (omitted in this study) and end-of-month next year. The questionnaire procedure is as follows:⁵ three to seven days before the end of the month the agency sends polls to the participants of the survey – mostly analysts at commercial banks operating in Poland; the number of banks participating in the survey varied from 12 to 21 in the period under review; the poll includes three tables to fill in – short- and long-term rates expectations, as well as central bank policy rate; the short-term results should be published on the last day of the month, the long-term about the ninth day of the following month, and the central bank rate at the end of the week preceding the MPC meeting; in practice however, most of the answers come to the agency towards the end of the first week of the month (i.e. after the publication of PMI indices, but before the Central Statistical Office of Poland announces other macroeconomic data), and are published subsequently.⁶ In step with the available poll horizons, we used mid-rates of FRA1x4 contracts as a market-implied one-month ahead expectation of the short-rate, and FRA13x16 for the longer-term expectation. In fact, since FRA13x16 contracts are not traded, we interpolated quotations of the FRA12x15 and FRA15x18 rates linearly to the required tenor. Thus, end-of-month readings of the above contracts represent FRA-implied expectations of the WIBOR 3M one-month and thirteen-month ahead. Figure 4 presents end-of-month survey-based expected WIBOR 3M and the corresponding FRA1x4 quotation, while Figure

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⁵ We would like to acknowledge Marcin Goettig of Thomson Reuters, Warsaw office, for providing us with the relevant datasheets and explanations on the questionnaire procedure.

⁶ The survey was suspended in December 2010 and in January 2011; we used linear interpolation (on a series of medians) to fill in the data for this period.

5 presents the end-of-month next year survey and market-based values. Table 2 and Table 3 present their basic descriptive characteristics.

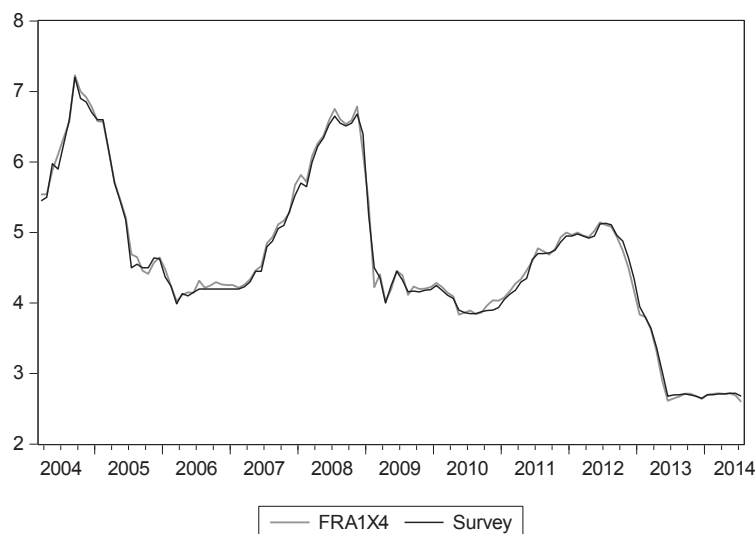


Figure 4. WIBOR 3M: short horizon market and survey expectations (%).
Source: Thomson Reuters.

Table 1. Descriptive statistics of the one-month ahead survey and market based expectations of WIBOR 3M.

	Mean	Median	Std. Dev.	Kurtosis	Skewness	Range	Min	Max	Auto-corr. coef.
<i>Survey</i>	4.62	4.45	1.11	2.69	0.24	4.55	2.65	7.20	99
<i>FRA 1x4</i>	4.64	4.45	1.13	2.69	0.21	4.63	2.60	7.23	99

Note: all statistics are quoted in percentage points.

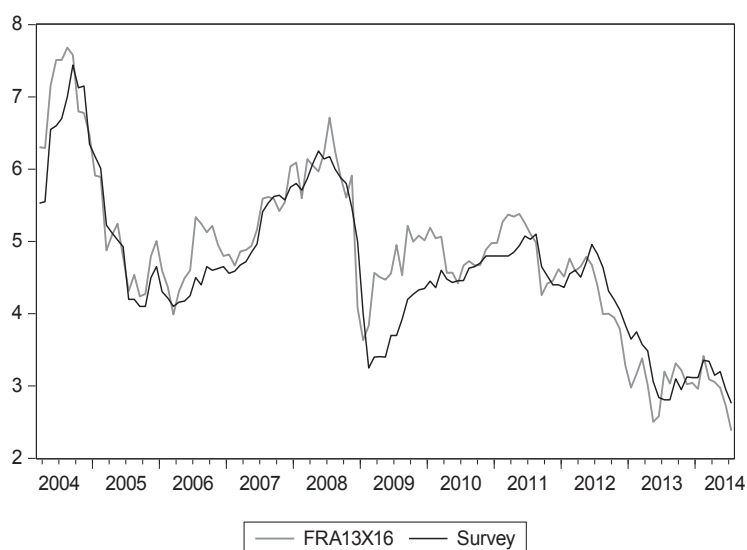


Figure 5. WIBOR 3M: long horizon market and survey expectations (%).
Source: Thomson Reuters.

Table 2. Descriptive statistics of the thirteen-month ahead survey and market based expectations of WIBOR 3M.

	Mean	Median	Std. Dev.	Kurtosis	Skewness	Range	Min	Max	Auto-corr. coef.
<i>Survey</i>	4.64	4.60	1.02	2.99	0.38	4.68	2.76	7.44	99
<i>FRA 13x16</i>	4.82	4.80	1.12	3.11	0.15	5.30	2.38	7.68	99

Note: all statistics are quoted in percentage points.

As explained in the introduction, we identify pure market expectations of the short-rate with the poll median, and the term premium with the FRA-median differential. We will thus call the one-month ahead term premium implied from the data in Figure 4 the shorter term premium (denoted tpS), while the term-premium implied from the data in Figure 5 thirteen months ahead – the longer term premium (denoted tpL). Thus, if the analysis allows one properly restricted cointegration relationship, the term premia can be obtained as error correction terms of the respective two-dimensional systems. The section below is devoted to this preliminary task.⁷

Cointegration analysis

First, consider data consisting of market FRA1x4 quotations and end-of-month survey-based expectations of the WIBOR 3M.

For both series, autocorrelation plots show slowly decaying values and partial correlation plots show significant values up to lag three. Since we are interested in modelling the autocorrelation structure of the series, this is the original motivation to use univariate autoregressive models in the present paper. Since the series do not possess any noticeable seasonal pattern, time trend or shift, the only deterministic term we included in the specification is a constant. To decide on the proper lag order for the models, we applied typical information criteria (Akaike, Hannan-Quinn, Schwarz), which clearly point to AR(3) and AR(2) models for FRA and median-expected WIBOR 3M respectively. The autocorrelation plots show slowly decaying values, although not that slowly as usual for unit root series. However, the Augmented Dickey-Fuller (ADF) unit root tests (Dickey & Fuller, 1979) cannot reject the unit root null, therefore we regard the series as nonstationary. The portmanteau

⁷ The results of all econometric procedures, including plots and tests, not presented in the text are available upon request from the authors.

and Ljung & Box tests (Ljung & Box, 1978) for residual autocorrelation with 16 lags prove insignificant, thus fitted models sufficiently filter out correlation from both series. The plots of standardized residuals occasionally exceed the conventional significance threshold (if any) in the first half of the sample and around the outbreak of the global financial crisis in 2008.

Table 3. ADF model estimation results for FRA1x4 quotations.

Variable	Const.	FRA_{t-1}	ΔFRA_{t-1}	ΔFRA_{t-2}
Coefficient	0.13	-0.03	0.34	0.23
t-statistic	1.64	-1.84	3.78	2.63

Note: asymptotic critical values are -3.43 (1%), -2.86 (5%), -2.57 (10%).

Table 4. ADF model estimation results for the one-month ahead expected WIBOR 3M.

Variable	Const.	$Survey_{t-1}$	$\Delta Survey_{t-1}$
Coefficient	0.10	0.02	0.50
t-statistic	1.34	-1.53	6.33

Note: asymptotic critical values are -3.43 (1%), -2.86 (5%), -2.57 (10%).

Moving on to determining the cointegration rank, we employed standard Johansen trace test (Johansen, 1995), again in a model with a constant as the sole deterministic term. The information criteria suggest one lagged difference should be used in the VEC representation. Bearing in mind univariate results and the fact that choosing too small an order can lead to size distortions for the tests, which is more harmful than selecting too large an order (implying merely a reduction in power), we decided to use two lagged differences instead. The results of the test given below, decisively point to one cointegration relationship. As in the univariate models, residual autocorrelation plots or tests for any remaining autocorrelation do not point to model adequacy problems and the standardized residuals very rarely hit the significance threshold.

Table 4. Johansen trace test for FRA1x4 and one-month ahead expectations of WIBOR 3M.

Cointegrating rank	Test statistics	Critical values			p-value
		90%	95%	99%	
r=0	38.99	17.98	20.16	24.69	0.000
r=1	4.35	7.60	9.14	12.53	0.375

In the next step we test whether the cointegrating vector, normalized to have a coefficient of one on the first variable, can be further restricted to have a coefficient equal to minus one on the second. If this proves right, the cointegration relationship can be interpreted as the term premium, i.e. as the difference between the FRA quotation and the median expected WIBOR 3M. The Wald test for beta restrictions using Johansen ML estimator has one degree of freedom and results in the test statistic amounting to 5.8, which translates into the p-value of 0.02. At the conventional 5% this would imply rejection of the desired restriction, but the evidence is not that strong since the p-value is still larger than 1%. Guided by the theory, therefore, we opt to maintain the hypothesis.

Standard cointegration analysis would proceed by re-estimating the restricted model, performing a model reduction and finally model adequacy tests, before moving to interpretation through Granger causality, impulse response analysis or forecast error variance decomposition. For the purpose of this paper it is, however, unnecessary, since our aim in this section was to prove the existence of the cointegration relationship, which could be interpreted as the term premium. The relationship will be further analysed on its own in the next section.

Consider now data consisting of market FRA13x16 quotations and survey-based expectations of the WIBOR 3M end-of-month next year.

Autocorrelation plots show slowly decaying values and partial correlation plots show significant values up to lag one for FRA and three for the survey. Again the only deterministic term that we included in the autoregression specification is a constant. Typical information criteria unanimously point to AR(1) and AR(3) models for FRA and the survey respectively. The ADF unit root tests cannot reject the unit root null, therefore we regard the series as nonstationary. The portmanteau and

Ljung & Box tests demonstrate that the fitted models sufficiently filter out correlation from both series and standardized residuals only occasionally exceed the conventional significance threshold.

Table 5. ADF model estimation results for FRA13x16 quotations

Variable	Const.	FRA _{t-1}
Coefficient	0.15	-0.04
t-statistic	1.08	-1.33

Note: asymptotic critical values are -3.43 (1%), -2.86 (5%), and -2.57 (10%).

Table 6. ADF model estimation results for the thirteen-month ahead expected WIBOR 3M

Variable	Const.	Survey _{t-1}	Δ Survey _{t-1}	Δ Survey _{t-2}
Coefficient	0.20	-0.05	0.25	0.26
t-statistic	1.92	-2.11	2.87	2.94

Note: asymptotic critical values are -3.43 (1%), -2.86 (5%), and -2.57 (10%).

The information criteria suggest a two lagged difference should be used in the VEC representation. The Johansen trace test (again in a model with a constant as the sole deterministic term), whose results are given below, actually points to no cointegration relationship. This is not desirable for the purpose of the paper. However, the p-value for the null of no cointegration just exceeds the 5% threshold by one percentage point. Bearing in mind the low power of cointegration tests and the result of the following section, where it is demonstrated that the FRA – median spread actually *is* stationary, we decided to move on under the hypothesis of one cointegration relationship. As in the univariate models, residual autocorrelation plots or tests for any remaining autocorrelation do not point to model adequacy problems and the standardized residuals very rarely hit the significance threshold.

Table 7. Johansen trace test for FRA13x16 and the thirteen-month ahead expectations of WIBOR 3M

Cointegrating rank	Test statistics	Critical values			p-value
		90%	95%	99%	
r=0	19.54	17.98	20.16	24.69	0.061
r=1	3.64	7.60	9.14	12.53	0.479

In the next step we test the relevant restriction on the cointegration space. The Wald test for beta using Johansen ML estimator results in the test statistic amounting to 1.7 which translates into the p-value of 0.19. We thus maintain the null hypothesis, which means that the longer term premium is the cointegration relationship in the VEC model.

Term premia stationarity and persistence

We are therefore justified to treat the term premia as new variables and analyse them on their own. First, we fit univariate models and check for stationarity. This can be considered as a robustness check of the analysis above, since the cointegration relationship, if it exists, should be stationary. Second, we answer the central question of the paper, i.e. we investigate which term premium is more persistent.

Recall that we dubbed the one-month ahead term premium the shorter term premium and we denote it by tpS , while the one-month next year term premium is called the longer term premium and denoted by tpL . Figure 5 below presents both premia, while Table 8 shows their basic descriptive statistics. The one-month ahead term premium is on average lower than the longer term premium, the difference being 14 bp. The visual inspection of the plot shows higher volatility of the longer term premium, tpL , compared to the shorter term premium tpS . It also suggests that while the tpS looks stationary, persistence in the tpL may be substantially higher, which likens it to a non-stationary process, its variance being bounded, however. The two are also correlated with each other at around 35%.

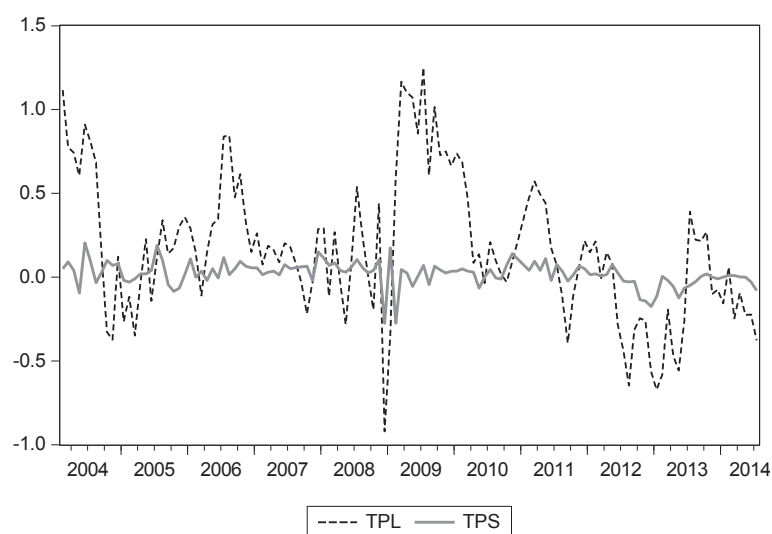


Figure 5. Term premia in the Polish money market (%).

Table 8. Descriptive statistics of the term premia.

	Mean	Median	Std. Dev.	Kurtosis	Skewness	Range	Min	Max	Auto-corr. Coeff.
<i>tpS</i>	0.02	0.03	0.07	6.28	-1.04	0.48	-0.28	0.21	14
<i>tpL</i>	0.17	0.15	0.42	2.97	0.29	2.17	-0.92	1.25	74

Note: all statistics are quoted in percentage points.

The analysis of autocorrelation and partial autocorrelation functions, ACF and PACF respectively, leads to the proposed autoregressive model for both term premia. The partial autocorrelation function of the *tpS* has a significant spike at the second lag, while the spike of the *tpL* is significant only at the first lag. Additionally, the ACF of the *tpL* shows a regular, decaying pattern. Thus, we expected autoregressive processes of rather low order.

The appropriate order of the autoregressive models was determined using information criteria such as Schwarz's, Akaike's and Hannan-Quinn's. The criteria unanimously suggest that the *tpS* follows an autoregressive process of order two, while the *tpL* follows a process of order one. Therefore, we specified AR(2) and AR(1) models for the *tpS* and the *tpL* correspondingly. The resulting estimated models for the *tpS* and the *tpL* with standard errors in parentheses, p-values in brackets and the coefficient of determination R^2 to the right are as follows:

$$(tpS) \quad y_t = 0.02 + 0.087y_{t-1} + 0.3y_{t-2} + \varepsilon_t, \quad R^2 = 11\%;$$

(0.010) (0.087) (0.087) [0.062] [0.317] [0.001]

$$(tpL) \quad y_t = 0.133 + 0.739y_{t-1} + \varepsilon_t, \quad R^2 = 57\%.$$

(0.095) (0.059) [0.163] [0.000]

The non-zero probabilities of the Breusch-Godfrey Lagrange multiplier tests and F-tests indicated no serial correlation in residuals for both models. The lagged correlations of the residuals in *tpS* and *tpL* models are close to zero except for some episodic insignificant fluctuations and one significant spike at lag 8 for the *tpL*; however, the Ljung-Box Q-statistics and corresponding non-zero probabilities agree with the previous tests in favour of a lack of serial correlation in the residuals of both autoregressions. Hence, there is no serious reason for concern about the adequacy of the models.

Turning to the issue of stationarity of the term premia, the modulus of the inverted roots of the AR(2) model are 0.59 and 0.51, while the inverted characteristic root of the AR(1) equation is 0.74; therefore, the estimated AR coefficients represent stable, stationary processes. The stationarity was tested formally and finally confirmed using the ADF and alternative (nonparametric) Phillips-Perron (PP) unit root tests, whose results are given in Table 9 below.

Table 9. Results of the ADF and PP unit root tests of the term premia.

Variable	Deterministic Terms	ADF				PP		
		Lags	Test value	5% critical value	p-value	Test value	5% critical value	p-value
<i>tpS</i>	<i>c</i>	2	-5.33	-2.89	0.0000	-9.67	-2.89	0.0000
	<i>c, t</i>	2	-5.81	-3.45	0.0000	-10.44	-3.45	0.0000
	<i>none</i>	2	-4.997	-1.94	0.0000	-9.04	-1.94	0.0000
<i>tpL</i>	<i>c</i>	1	-4.45	-2.89	0.0004	-4.45	-2.89	0.0004
	<i>c, t</i>	1	-4.61	-3.45	0.0016	-4.61	-3.45	0.0016
	<i>none</i>	1	-4.27	-1.94	0.0000	-4.20	-1.94	0.0000

The calculated results of the measures of persistence introduced in Section 4 are presented in Table 10 below. One evident observation is that all the measures demonstrate higher persistence in the case of the longer term premium, *tpL*. This

perfect consensus is reassuring and seems to be due to the relatively simple dynamics of the series – autoregressions of very low orders. Thus, the *tpL* is not only more volatile, but also more persistent, than the *tpS*.

Table 10. Persistence measures of the term premia.

	ρ	<i>CIR</i>	γ	<i>P</i>	<i>h</i>
<i>tpS</i>	0.39	1.63	0.67	0.03	1
<i>tpL</i>	0.74	3.83	0.72	0.16	3

In particular, both the sum of autoregressive coefficients and the cumulative impulse response measures are roughly twice as high in the case of the longer term premium as for the shorter one. The unconditional probabilities of not crossing the mean in a given period are higher than 50% in both cases pointing to a positive persistence, and the probability is higher for *tpL* by around 5 percentage points. The random-walk component responsible for persistence appears rather unimportant in the shorter term premium, while it is responsible for up to 16% of variance of the *tpL*. Finally, the shorter term premium halves the distance to its equilibrium in just one month after the deviation has occurred, while the longer term premium needs the whole quarter to do it.

6. Conclusions

This paper focuses on the Polish money market term premia over a ten-year period. To identify the premia, we applied the method of obtaining the expected interest rate by relying on survey data, which is a relatively recent approach for emerging countries, and entirely novel for Poland. As a forward rate, we chose the forward rate agreements, whose market is a very liquid segment of the financial market in Poland. We presented the economics of the FRA instrument and described the Polish FRA market as well. Then we carried out econometric analysis of the shorter and longer term premia during which it was found that both time series are stationary autoregressive processes of low order.

A significant part of our work deals with measures of the term premia persistence. We applied five measures of persistence, all well-performing according to the reviewed literature. Our findings unambiguously indicate higher persistence in the case of the longer term premium. Moreover, the longer term premium is more volatile than the shorter one. The presence of persistence points to the rejection of the EMH at the short end of the Polish interbank yield curve.

As we highlighted at the beginning, this study is only the first step in the investigation of the Polish term premium in the money market. In the next paper we plan to improve the fit of the models to solve some model adequacy problems that we encountered and to take account of a possible break in the middle of the sample due to the outbreak of the 2008 crisis. Another interesting extension of the current work that we also aim to pursue involves searching for economic determinants of the term premia investigating its forecasting power in financial markets.

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