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Do Inflation Expectations Surveys Yield Macroeconomically Relevant Information?

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1. Introduction

Money market prices have an important role in the transmission of monetary policy. The prices are influenced not only by the monetary policy decisions but also by the beliefs of market makers. According to the Czech National Bank (CNB or Bank), the surveyed inflation expectations do not influence policy decisions. They are only collected for informative purposes. The Bank's own inflation forecasts dominate the private forecast in policy decision making. It is believed that the Bank has more and better information about the economy and its future prospects than private agents. In the adaptive learning theory – see e.g. (Evans – Honkapohja, 2003) – which is likely to apply to the Czech economy, if the expectations of the central bank and private agents differ, and, at the same time, the private agents' expectations form market prices, it might be desirable for the central bank, in order to meet its objectives of inflation and output gap variability minimization, to set its instruments given the private expectations.

The Czech National Bank runs monthly surveys on the inflation expectations of representative institutions operating on the Czech financial market. In this paper we analyze, from the macroeconomic perspective (which will be explained later), the relevance of these expectations for the monetary policy, i.e. their role in the formation of prices in the Czech Interbank money market. The question we ask is: *To what extent do the surveyed expectations correspond to the true market expectations?* Or put differently, *do the surveyed expectations include any monetary-policy relevant information?*

This paper builds on the following four observations which all together provide grounds for assessing the relevance of private forecasts for the mo-

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FIGURE 1 PRIBOR 1Y and Discount Rate

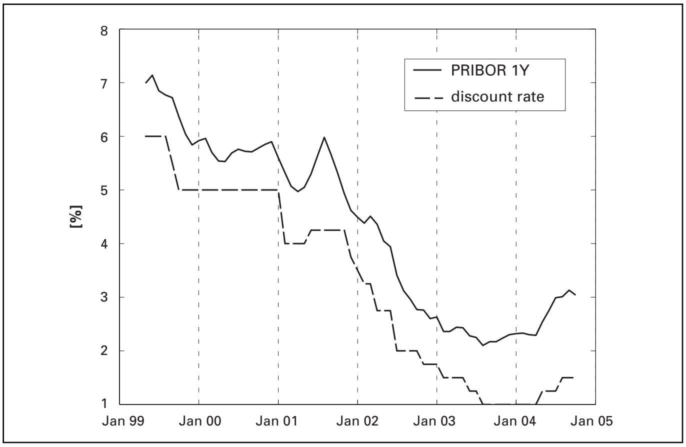
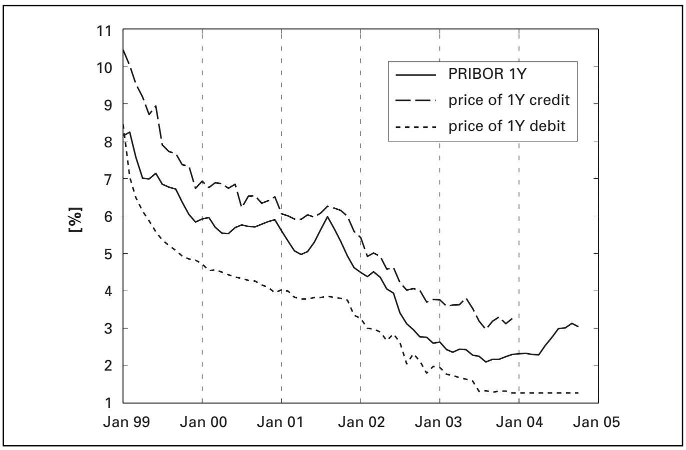


FIGURE 2 The Interbank Money-market Price, and Credit and Deposit Market Prices



netary policy. When interpreting the results in this paper, one should bear in mind that:

1. the CNB sets its policy instrument given its own inflation forecasts (see CNB's Inflation Reports);
2. the CNB's inflation forecasts are conducted on a lower frequency (quarterly) than the financial-analysts-inflation-forecast surveys which are formed on a monthly basis;
3. the CNB's discount rate (policy instrument) influences the mean values of market interest rates (in this paper we focus on 1Y PRIBOR) (*Figure 1*). The policy instrument adjusts on a lower frequency than money market prices;
4. the money-market price influences the mean value of credit-market prices (*Figure 2*).

FIGURE 3 Comparison of Inflation Expectations and Actual Inflation

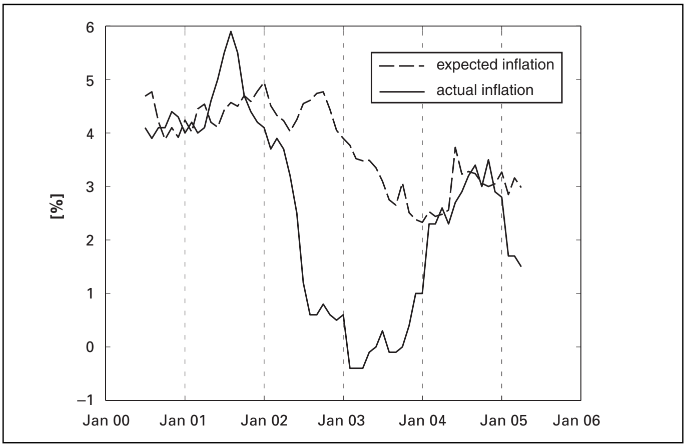
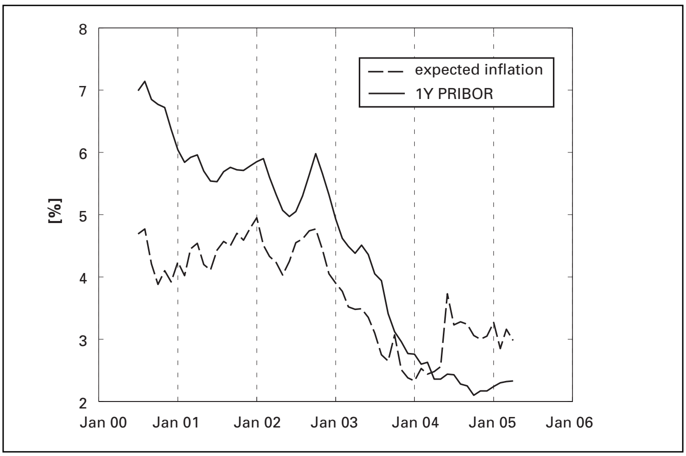


FIGURE 4 Comparison of Inflation Expectations and PRIBOR 1Y



A simple plot of the surveyed expectations and the actual inflation rate in *Figure 3* reveals that the expectations exhibit a tendency for overshooting and inefficiency. What is remarkable, however, *Figure 4* shows very similar behavior of the money-market price (PRIBOR – Prague Interbank Offer Rate) and of the surveyed expectations. A very similar development of PRIBOR and private forecasts is significant till March 2004. Then a structural break occurred. After an upward correction for the impact of the tax reform and of the expected growth in world oil prices due to the second war in Iraq, it seems the nominal interest rate and expectations follow a similar trajectory again, although on a different level. The *ex ante* and also *ex post* real return on the Interbank market are negative now.

Given these observations, our interest is to analyze whether the surveyed data has any information content, i.e., whether the financial analysts' expectations have any impact on the market data. We may observe an interesting paradox. In *Figure 3*, the expectations do not have any predictive power for the actual inflation rate.[1] However, in *Figure 4*, it seems the expectations are reflected in the money market price. It is true that the money market interest rate is, in general, given by a monetary policy rule (see the observation 1 and 3 above). Having its own forecasts, the central bank sets the magnitude of its instruments. This is done however on a low (quarterly) frequency. Looking at Figure 4, the variability of private forecasts is projected in the variance of the interest rate. If the Interbank interest rate transmits in the credit market rates (see observation 4), one should suspect that the private expectations has economic relevance even though they do not have a predictive power for actual inflation. The paper tests this in a formal way.

Forming expectations is an essential part of any optimal decision making of any individual. Macroeconomic theory builds on the rational expectation hypothesis (REH). Certainly, because agents are heterogenous (they face different information utilization costs, they have different information sets available, etc.), expectations about the future differ. For the REH it is important that the expectations are on average (of all individuals) unbiased and efficient, i.e., on the aggregate level the expectations correspond to reality with only minor and random differences.

Recently Mankiw and Wolfers (2003) provided empirical evidence (for the American economy) that inflation expectations differ for different groups of agents (households, firms, academics, professional forecasters, etc.), which they explain by a sticky information model. According to their model, the disagreement in expectations varies with the state of the economy. However, over all economy states financial analysts form the most efficient and consistent inflation forecasts from all the groups Mankiw and Wolfers study. It is because the financial analysts have the best market-related information. If we assume that this applies to the Czech conditions as well, it calls into question why the forecasts in Figure 3 are inefficient and inconsistent and why they still have predictive power on the price on the Interbank market.

In this paper we only deal with the expectations of a single group of people, financial analysts. As such, one would think that what we observe in Figure 3 is irrelevant for judging the applicability of the assumptions of the REH which require we survey a representative sample over *all* economic agents (households, firms, financial intermediators, etc.). This paper is based on the fact that the Interbank money market is dominant in the Czech financial system, unlike the US market. The CNB surveys a representative sample of financial analysts who represent the major market-makers on the Czech Interbank market. The price the market-makers create (PRIBOR) strongly influences other financial market prices (credit, swap market prices, etc.). Thus, if the financial analysts' expectations play a role in the determination of PRIBOR, then judging their properties and the usability for macroeconomic analysis is appropriate.

[1] They rather seem to be consistent with the CNB's inflation target.

A careful reader will have already noticed that we view the money market from the macroeconomic (modeling) perspective, and that it will be abstracted from the real mechanism of how the market functions, i.e., how the market price is actually formed. The Fisher rule is the result of any macroeconomic model based on the first principles. Under certain assumptions, the Fisher rule can also be view as an optimal monetary policy rule – see e.g. (Woodford, 2003, Ch. 2). As such it will be a cornerstone of our testing metodology, similar to Fama and Gibbons (1982), Mishkin (1990a) or Mishkin (1990b).

The rest of the paper is structured as follows. The methodology is formulated in the next section 2. The data set and estimation results with their interpretation follow in sections 3 and 4, respectively. Section 5 concludes with a discussion of the results.

2. The Methodology

The methodology developed in this section is structured so that the stated question of interest "*To what extent do the surveyed expectations correspond to the true market expectations?*" is addressed. To this purpose the null hypothesis is formulated as: "*The market expectations coincide with the surveyed expectations.*"

It ought to be pointed out that in the methodology developed below the expectations' rationality, as assumed by the REH, is not questioned. Instead, the assumption is weakened and the expectations may be biased and inefficient.

Testing H_0, similar to Fama and Gibbons (1982), Mishkin (1990a) and Mishkin (1990b), the Fisher rule is assumed as the true pricing rule for the money market, i.e.,

$$i_t^m = \Pi_{t,m}^e + r_t^m + \rho_t \tag{1}$$

where i_t^m is the nominal interest rate valid from the beginning of period t to $t + m$, $\Pi_{t,m}^e$ are the true market expectations formed at the beginning of period t for the $t + m$ time horizon, and r_t^m is the corresponding *ex ante* real interest rate. The Fisher rule further includes the risk premium ρ_t, which is assumed to follow a stationary process uncorrelated with expected inflation and the real interest rate.

In a standard way, let us assume that the inflation expectations are based on information Ω available at time t. Furthermore, following assumptions common to the literature, let us assume that the market inflation expectations at time t for the $t + m$ horizon are equal to the actual inflation rate π_{t+m}, but are subject to a random error ε_{t+m}. Formally written,

$$\Pi_{t,m}^e = E(\pi_{t+m} \mid \Omega_t) = \pi_{t+m} + \varepsilon_{t+m} \tag{2}$$

Under the REH, it holds that ε_{t+m} is iid with zero mean and finite variance. If the mean is non-zero, then the expectations are biased. If the expectation error follows a stationary process with zero mean, the expectations are so-called weakly rational. For the testing procedure developed below, no particular form of the market expectations is prescribed, hence they can be both rational or weakly rational, i.e. ε_{t+m} can be iid or have an AR structure, respectively. Recalling the non-zero auto-covariance structure causes only inefficiency of expectations but does not affect their consistency.

As the objective is to find the relation of the surveyed expectations to the market ones, it is convenient and sufficient to keep ε_{t+m} in a nearly unspecified form.

Now substituting (2) in (1) yields

$$i_t^m = \pi_{t+m} + \varepsilon_{t+m} + r_t^m + \rho_t \tag{3}$$

Depending on the correlation of expected inflation, $(\pi_{t+m} + \varepsilon_{t+m})$, and *ex ante* real interest rate, r_t^m, as argued in (Mishkin, 1990b), the relation between nominal interest rate and expected inflation does not need to be one-to-one. Hence, we augment (3) into a more general form

$$i_t^m = \phi_1 (\pi_{t+m} + \varepsilon_{t+m}) + r_t^m + \rho_t \tag{3a}$$

The Fisher rule where $\phi_1 = 1$ counts only for a special case.

Now adding and subtracting the mean values of the real interest rate, unanticipated inflation and risk premium $(\phi_1 E(\varepsilon_{t+m}), E(r_t), E(\rho_t))$, (3a) becomes

$$i_t = \phi_0 + \phi_1 \pi_{t+m} + \nu_t \tag{4}$$

where

$$\phi_0 = \phi_1 E(\varepsilon_{t+m}) + E(r_t) + E(\rho_t)$$

and

$$\begin{aligned}\nu_t = \phi_1 \varepsilon_{t+m} + r_t + \rho_t - [\phi_1 E(\varepsilon_{t+m}) + E(r_t) + E(\rho_t)] = \\ = \phi_1 (\Pi_{t,m}^e - \pi_{t+m}) + r_t + \rho_t - \phi_0\end{aligned}$$

The constant term ϕ_0 will be in general different from zero. As discussed above, no specific assumptions are imposed on ε_{t+m}, and thus in general $E(\varepsilon_{t+m})$ may be non-zero. Similar can be said about the mean values of the *ex ante* real interest rate, $E(r_t)$, and the risk premium, $E(\rho_t)$, which both can be expected to be of positive magnitudes.

For ν_t in formulation (4), one is not able to separate the real interest rate and risk premium from the expectation error. It is impossible to distinguish what part of the variation of ν_t accounts for the unanticipated inflation and what part accounts for the ex ante real interest rate and risk premium variability. One is not able to extract the error in the market expectations from ν_t, and to learn the expectations' exact form.[2] The testing procedure avoids this limitation. The empirical test is built solely on the *comparison* of expectations processes where only one of them has to be directly observable.

[2] On this fact Cambell, Lo and MacKinlay (1997, section 1.5.2.) build their argument that the rational expectations hypothesis is not testable. It can be only assumed.

The null hypothesis assumes a correspondence between the surveyed expectations (observable process) and the market expectations (unobservable process).

The error term ν_t in expression (4) can be written for the market expectations and survey expectations separately:

$$\nu_t = \phi_1 (\Pi^e_{t,12} - \pi_{t+12}) + r_t + \rho_t - \phi_0 \tag{5}$$

and

$$\nu\nu_t = \psi_1 (\pi^e_{t+12} - \pi_{t+12}) + rr_t + \rho\rho_t - \psi_0 \tag{6}$$

where ψ_0 and ψ_1 have the same interpretation as ϕ_0 and ϕ_1, respectively, but as the surveyed expectations $\pi^e_{t,12}$ are introduced a different denotation is chosen. From here the idea of the test is straightforward.

The null hypothesis is

$$H_0: \qquad \pi^e_{t,12} = \Pi^e_{t,12}$$

where again $\pi^e_{t,12}$ denotes the surveyed inflation expectations formed at the beginning of period t for the time horizon till $t + 12$, and $\Pi^e_{t,12}$ are the true market expectations for the same period of time. We can see that under H_0, the error terms (5) and (6) must coincide, i.e. $\psi_0 = \phi_0$ and $\psi_1 = \phi_1$, and including the unobservables, i.e., $r_t = rr_t$ and $\rho_t = \rho\rho_t$.

Substituting $\nu\nu_t$ instead of ν_t in (4) and rearranging, one obtains

$$i_t = (\phi_0 - \psi_0) + (\phi_1 - \psi_1)\, \pi_{t,12} + \psi_1 \pi^e_{t,12} + rr_t + \rho\rho_t \tag{7}$$

For the testing purposes, we set $\phi_0 - \psi_0 = a_r$, $\phi_1 - \psi_1 = a$, $\psi_1 = b$ to obtain the final testing formula

$$i_t = a_r + a\pi_{t+12} + b\pi^e_{t,12} + u_t \tag{8}$$

Because the *ex ante* real interest rate and risk premium are unobservable, they are put as the error term u_t in (8), $u_t = r_t + \rho_t$. If the above assumptions and following requirements hold, then one may conclude that $\pi^e_{t,12} = \Pi^e_{t,12}$. (i) if $\psi_0 = \phi_0$ and $\psi_1 = \phi_1$, then it must be that $a_r = a = 0$ and (ii) b must be statistically significant and positive. From (i), we see that testing H_0 on (8), we perform a joint hypothesis test. Firstly, we test for the equality of surveyed and market expectations. And secondly, we test for the equality of *ex ante* real interest rates and risk premiums.

When performing the test, one ought to be aware of the three major weaknesses the test suffers from. The first concerns the *character of disturbances* in (8).

The disturbances in (8) contain the variable part of ex ante real interest rate. If the real interest rate is a function of productivity, it may be the case that the disturbance term follows a non-stationary process. In general, when residuals contain a unit root, one ends up with a spurious regression. Given this possibility, an essential part of the methodology is to test the disturbances for stationarity. The augmented Dickey-Fuller (ADF) test and the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test are employed. Indeed, if a unit root is identified, the test cannot be used.[3]

If the disturbances in (8) do not follow a unit-root process they are likely to have an autocorrelation structure. In this case the test can still be used, one just has to correct for the autocorrelation effect on the parameters' standard error estimates. The parameters' unbiasedness and consistency property remain unaffected otherwise.

Further in equation (3a) it is stated that the nominal interest rate and expected inflation do not need to move in a one-to-one relation. Following the analysis by Mishkin (1990b), this may arise if there is correlation between expected inflation and ex ante real interest rate. Since, the real interest rate is in the error term, OLS estimates are inconsistent and the instrumental variable approach must be employed instead. This is the method used here.

The second problem one should be aware of is the *problem of multi--collinearity*. The occurrence of multi-collinearity is very likely, which follows from the nature of the formation of expectations. The closer the expectations are to the REH, the more severe the problem is, since both π_{t+12} and $\pi^e_{t,12}$ are regressors.

The third concern is the *problem of joint hypothesis*. The problem is that the methodology crucially relies on the assumption that the Fisher rule holds. Thus if it happens the null hypothesis is rejected, one cannot be certain whether it is because the market expectations are truly different from the surveyed ones or because the Fisher rule is an invalid assumption. Indeed, this is common to all tests of this type.[4] However, the problem is lessened here. As mentioned in the introductory section, the focus on the problem of inflation expectations is from a macroeconomic perspective, and there the Fisher rule is a standard way to capture the interest rate behavior.

In summary, to test H_0, we estimate (8) and test whether the parameters a_r, a, and b are statistically significant. If the parameters a_r and a are insignificant and the parameter b is significant then H_0 cannot be rejected. If H_0 cannot be rejected, the surveyed inflation expectations are likely to coincide with the market expectations. Accounting for the test weaknesses, an essential part of the test has to be (i) the test of residuals stationarity and the consequent adjustment of critical values, and (ii) a check of the multi-collinearity magnitude.

3. The Data

The data set consists of the monthly data of the Czech Interbank money market. The source of the data is the Czech National Bank and the Czech Statistical Office. For the testing purposes three series of monthly data are employed: (i) the monthly average of nominal one-year Interbank interest rate (Prague InterBank Offer Rate, PRIBOR 1Y), (ii) year-to-year CPI inflation, and (iii) year-to-year expected change in the CPI. The expectations have been collected at the beginning of each month. The details of the sur-

[3] The Bohm-Bawereck hypothesis may offer a solution. As the latent variable for the *ex ante* real interest rate, the real GDP growth can be used. Following the same idea as with expected inflation, the real GDP growth becomes one of the explanatory variables in (8), and the unanticipated part of the GDP growth enters the error term. In a standard economic environment, the unanticipated real growth ought to be stationary and so does the whole error term in (8).

[4] Cf. (Cambell – Lo – MacKinlay, 1997, section 1.5.2).

vey can be found at *http://www.cnb.cz*. The data range is from 1999:05 to 2005:05. The whole data set can be obtained upon a request from the author.

Before moving further, let us discuss the data timing which is crucial for the test. A period t denotes a month. PRIBOR 1Y is an average of daily rates within the period t. CPI is a measure of the price level within a given period t. The data to construct the CPI are typically collected in the second week of each month. We can roughly think of it as of a monthly price level average. The data itself is published in the beginning of the following month. Thus when computing the actual inflation rate one must be careful about the data timing. For our purposes, the year-to-year inflation rate at time $t + 12$ is the relative change in CPI between period $t - 1$ and $t + 11$. Because the information set available to the agents at time t contains only the CPI of $t - 1$ as the latest information about the actual price level, i.e., $\Omega_t = \{CPI_{t-1}, CPI_{t-2},...\}$, then when comparing $\pi^e_{t,12}$ with actual inflation over this period, (π_{t+12} as denoted in equation (8)), the actual inflation rate is computed as $\pi_{t+12} = 1 - (CPI_{t+11}/CPI_{t-1})$.

4. The Results

First let us draw attention to the critical values of t-tests employed here. In the results below, a very low value of Durbin-Watson statistics is found which indicates either the presence of a unit root in the residuals or their strong positive autocorrelation. To test for the former, the augmented Dickey-Fuller and Kwiatkowski-Phillips-Schmidt-Shin tests are used. Since the results, summarized in *Appendix B*, do not suggest statistically significant evidence for the presence of the unit root, we argue in favor of a strong positive autocorrelation in the residuals. Because of the autocorrelation, the parameters' standard errors are biased downwards, and the pivotal statistics of t-tests are biased upwards. By simulating new critical values, the effect of autocorrelation on the test results is eliminated. The critical values reported below are not the standard ones but those adjusted for the autocorrelation effect. Details on the simulation are presented in *Appendix A*.

The next reason for simulating the t-test's critical values, instead of using the Student's distribution, is also the presence of multi-collinearity. *Appendix C* investigates this issue more closely. Because the variance inflation factor (VIF), the measure of the magnitude of collinearity, is 1.52, and a critical value is 10, it is concluded that multi-collinearity is not a significant problem. Despite this however, the parameters' standard errors are affected and simulating critical values partially accounts for it.

Finally, the estimation results of equation (8) are reported in *Table 1* and *Table 2*. The test is performed for two data samples. The first sample ranges from May 1999 to March 2003.[5] The second one ranges from May 1999 to May 2004. Two different samples are considered because of the structural

[5] The dating is from the expectations surveys perspective. For instance, May 1999 stands for the month when expectations for the period May 1999–May 2000 were formed.

TABLE 1 $i_t = a_r + a\pi_{t+12} + b\pi^e_{t,12} + \varepsilon_t$, 1999:05–2003:03

Parameter	Estimate	*t*-stat.	Critical value 5 %	DW	$\bar{R}^2$	No. of obs.
a_r	0.2094	0.4008	4.5052	0.32	0.79	43
a	0.1445	2.8471	4.3978			
b	1.1406	7.7296	4.6014			

TABLE 2 $i_t = a_r + a\pi_{t+12} + b\pi^e_{t,12} + \varepsilon_t$, 1999:05–2004:05

Parameter	Estimate	*t*-stat.	Critical value 5 %	DW	$\bar{R}^2$	No. of obs.
a_r	–0.5230	–1.1005	4.5052	0.62	0.84	58
$a_{r,\ dummy}$	–1.6769	–6.8732	4.5052			
a	0.0853	1.5851	4.3978			
b	1.3599	10.1255	4.6014			

break which occurred in March 2003. The jump correction in expectations, accounting for 1.17% increase in expected inflation, is explained by the expected growth of oil prices due to the beginning of the second war in Iraq, and an expected tax reform. Having the two samples, we control for the structural change effect on the results.

For the sample excluding the structural break (Table 1), the null hypothesis that the market expectations are equal to the surveyed expectations cannot be rejected at the standard level of significance. The parameters a_r and a are not significantly different from zero, while the parameter b is found to be significantly different from zero. The goodness of fit is about 80 % (without accounting for autocorrelation in residuals) which suggests the joint hypothesis problem is not binding.

Similar results are obtained for the complete data sample. Here, equation (8) is expanded for a dummy variable to control for the structural break. In Table 2 we see again the parameters a_r and a are not significantly different from zero, while the parameter b is significant and positive. The dummy variable parameter $a_{r,dummy}$ is found significant. Its value of –1.67 can be accounted to the change in the mean value of unanticipated inflation $E(\varepsilon_{t+m})$ which constitutes the constant term in (8) and which has shifted in March 2003 by about 1.2 %. The negative value of $a_{r,\ dummy}$ indicates, that the impact of the tax reform and other exogenous shocks was not anticipated in the market expectations and thus not reflected in the market price (or it was mitigated by other factors). Even though the requirements to accept the null hypothesis are not met here, because the constant terms of unanticipated inflation differ, we can conclude at least the variable parts of (surveyed and market) expectations are very similar. Parameters a and b have similar values and character for both data samples considered. We might conclude that the surveyed expectations are very close to the market expectations also on the full length sample. The variable components of expectations are still close to each other and the variable components are the ones most interesting for any analysis.

Certainly, it is a positive finding the surveyed expectations contain important information for macroeconomists and that they are not only some senseless numbers. They may be taken seriously when assessing future macroeconomic developments. Of course, if the test results also hold in the future.

Now abstracting from the macroeconomic view, the results seem somewhat surprising. The surveyed expectations are formed by financial analysts which do not have any direct connection to the market makers in order to influence their actions. In practice, it is more than likely that the dealers do not pay any attention to the analysts' inflation expectations, or rather the dealers do not have any inflation expectations at all. Their particular objective, as professionals, is to maximize their profit and their actions mostly have speculative motives. Despite this, the financial analysts' expectations get included in the market price formed by the market makers. Here the surveyed expectations have such an explanatory power on the market price that they can be viewed to be close to the market expectations character.

In the rest of the text let us offer a possible story how the financial forecasts' may get transmitted in the market price. The story ought to be taken as a motivation for further discussion and elaborations on this topic.

The character of the Czech Interbank money market and the credit market can offer one possible explanation of the above observations. The credit market is the main channel used to transmit the capital to the economy. Banks (lenders) are the market price setters, and borrowers are price-takers. The price on the credit market is derived from the price on the Interbank money market which, in contrast to the credit market, may be considered a competitive one.

For banks as the major lenders of capital, it is profitable to bias their inflation forecasts (expectations) upwards. If we believe that bankers take into account a nominal depreciation of money when forming the credit price, overshooting these expectations increases their ex post real revenues. Given that the borrowers are price-takers, overshooting is accepted. Let us assume that a bank on the Czech credit market sets the one-year nominal interest rate on credit so that it is composed of an individual PRIBOR 1Y estimate plus a risk premium and a profit margin. The individual bank's PRIBOR 1Y estimate is composed of a required minimum ex ante real return plus an expected nominal depreciation (expected inflation). Having a price on the credit market, the lenders face a possible lack or excess of loanable funds (deposits from clients). To utilize them, they are motivated to enter the Interbank market and trade them. Under the assumption that the pricing rule is the same for all banks in the credit market, there will be only a moderate correction in the market PRIBOR in order to make the market clear. The new market PRIBOR is recursively reflected in the price on the credit market.

In this story, the money-market dealers do not necessarily need to know the inflation expectations the market price includes or even how the credit price is formed. They are only "endowed" with an excess or deficit of money which they trade. The deficits or excesses of loanable funds are determined by the credit market, which is exogenous to the Interbank market.

5. Conclusion

In this paper we find that surveyed inflation expectations, even though they do not have a predictive power for actual inflation, they have a predictive power for the interest rate and that they do not statistically differ from the market expectations. This leads us to the conclusion that the measurements of financial market inflation expectations conducted by the Czech National Bank yield macroeconomically relevant data and they can be taken seriously when assessing future macroeconomic developments.

The analysis is conducted from the macroeconomic modeling perspective and the choice of testing hypothesis and the data set serves this purpose. The results question the macroeconomic modeling standards and analysis of optimal monetary policy conduct in the Czech economy. They should motivate a discussion about modeling standards of private agents' expectations in macroeconomic models. So far it is assumed private agents have perfect knowledge and form the so called model based expectations (rational and complete-knowledge expectations). The analysis indicates, however, this does not need to be the case of the Czech economy. This may have consequences for the optimal conduct of monetary policy. As Orphanides and Williams (2003), and Evans and Honkapohja (2003) show, optimal monetary policy differs for an economy where the agents have full knowledge about the economic structure, and for an economy where agents face imperfect knowledge. What is found optimal in the rational and perfect knowledge world is not optimal in the imperfect knowledge world, and *vice versa*. Macroeconomists and policy-makers ought to be aware of this fact.

Nonetheless, the research has produced more questions than answers. The findings call for a study on why the Czech Interbank money market anticipates inefficient expectations, and why there is no arbitrage incentive to improve this. Given that the market has a unique position in the financial system, in general, it is not possible that the results found here would hold in the long run. Hence, although we make some headway towards explaining how the inefficient inflation forecasts are transmitted into the market price, a serious attempt to find the final answer should still be made. Having a better understanding of how the market functions, we are in a better position to improve our economic models and to suggest optimal policies.

APPENDIX A

In this appendix, the methodology for obtaining the critical values reported in the text is outlined. The methodology relies on Monte Carlo experiments. Mishkin (1990b) was the motivation for this approach.

The methodology can be summarized in the few following steps:

1. Analyze the time series of model variables, i.e., PRIBOR 1Y, year-to-year inflation, and inflation expectations, on the unit root.
2. Apply the Box-Jenkins methodology on the data. The outcome ought to be an ARIMA(p,d,q) model with the best fit possible.
3. Simulate the estimated models from the previous step.
4. Using the simulated time series, estimate equation (8), and save the results on the *t*-tests.
5. Repeat steps 3 and 4 10,000 times.
6. From step 5 construct a new distribution for critical values.

The simulated critical values are reported in *Table 3*.

TABLE 3 The Simulated Critical Values

Parameter	25 %	10 %	5 %	1 %
a_r	1.7339	3.6738	4.5052	7.2765
a	1.7232	3.4252	4.3978	6.8293
b	1.8041	3.4825	4.6014	6.8393

APPENDIX B

The regression residuals from (8) are tested for a unit root here. Testing for a unit root is crucial for the regression results, because residuals are partially estimates of real interest rate, and there might be an economic reason to believe that a unit root is present. If it is so, then the results are spurious and non-usable.[6] To test for the unit root, the augmented Dickey-Fuller (ADF) and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) tests are employed. Each of them test for a unit root but from a different view. The null hypothesis of the ADF test is that a *time series is not stationary* while the null hypothesis of the KPSS test is that a *time series is stationary*. Applying both tests provides a more complex picture than using only one of them.

The results for the ADF test are summarized in *Table 4.* Following Enders (1995), the ADF test is based on testing H_0: $\gamma = 0$ in

$$\Delta\hat{\varepsilon}_t = \gamma\hat{\varepsilon}_{t-1} + \sum_{i=2}^{p}\beta_i\Delta\hat{\varepsilon}_{t-i+1} + \nu_t$$

where $\hat{\varepsilon}_t$ are the residuals from estimating (8).

TABLE 4 Augmented Dickey-Fuller Test Results

Parameter	Estimate	*t*-test	*t*-crit.	*F*-test (*p*-value)	*Q*-test (crit.val.)
γ	–0.2192	–7.5765	–1.95	5.65 (0.02)	17.28 (25.70)
β_1	–0.0115	–0.1414	–1.69		
β_2	–0.1728	–2.3414	–1.69		

[6] For details on spurious regression, refer to Granger and Newbold (1978) who experimentally demonstrated the consequences of unit root on regression results, and to Phillips (1986) who formalized their results.

The critical values in Table 4 are for the 5% level of significance. The null hypotheses of residuals being non-stationary is rejected ($\gamma \neq 0$).

The KPSS test results are summarized in *Table 5*. The test statistics are computed for the lag truncation parameter, l, from 0 to 8. As argued by Kwiatkowski et al. (1992), for $l = 8$ the test has the largest power. Including $l = 0$, the test also accounts for autocorrelation. The critical value for the test at the 5% significance level is 0.463.

TABLE 5 KPSS Test Results

l	0	1	2	3	4	5	6	7	8
$\hat{\varepsilon}_t$	**1.10**	**0.62**	**0.47**	0.38	0.32	0.29	0.26	0.25	0.24

The results from the KPSS test suggest that the regression residuals from estimating (8) are stationary. For $l = 2$ the test result is on the margin of statistical significance. Putting the result together with the ADF test, it may be concluded that the residuals do not contain a unit root and follow a stationary process. Consequently, the parameter estimates reported above are unbiased and consistent, although inefficient.

APPENDIX C

In this section the collinearity issue is addressed. First, a formal test for the presence of collinearity is performed. At the same time the test is also a test of the rational expectations hypothesis. Second, the effect of collinearity on parameters' estimates and their standard errors is quantitatively analyzed.

The Test of Collinearity

The test is standard to the econometric literature. It is based on a *variance inflation factor* estimation. Because it tests for the linear relationship between π_{t+12} and $\pi^e_{t,12}$, the test is at the same time a test of the REH.

The REH is usually tested on the following inflation-prediction equation – see e.g. (Bakhshi – Yates, 1998):

$$\pi_{t+12} = c_1 + c_2\pi^e_{t+12} + \nu_{t+12}$$

If the REH holds, parameter c_1 is zero, c_2 is equal to 1, and ν_{t+12} is an iid process with zero mean and finite variance. In this case the two variables π_{t+12} and π^e_{t+12} are evidently collinear.

When the inflation-prediction equation is estimated, $\bar{R}^2$ is used to evaluate collinearity. To this purpose a variance inflation factor (*VIF*) is computed:

$$VIF = \frac{1}{1-\bar{R}^2}$$

Usually, we face a problem of collinearity if $VIF > 10$.

The estimation results of the inflation-prediction equation are summarized in *Table 6*.

First, we can see that $c_1 \neq 0$, $c_2 \neq 1$, and the residuals are positively auto-correlated. As a consequence, the REH cannot be accepted. Second, the *VIF* is 1.5 which is far from the value where collinearity causes estimation problems.

TABLE 6 $\pi_{t+12} = c_1 + c_2\pi^e_{t,12} + \nu_{t+12}$

Parameter	Estimate	*t*-stat.	Critical value 5 %	DW	$\bar{R}^2$	No.of obs.
c_1	–4.3353	–2.9835	2.5160	0.16	0.34	44
c_2	1.7051	4.8067	2.6748			

Note: The critical values are simulated similarly as in Appendix A.

The Quantitative Assessment of the Collinearity Effect on Estimation

Let us derive the estimates $\{\hat{a}_r, \hat{a}, \hat{b}\}$ of equation (8):

$$i_t = a_r + a\pi_{t+12} + b\pi^e_{t,12} + \varepsilon_t$$

To find the estimates of the parameters, the ordinary least-square criteria is used:

$$\sum_{t=1}^{N} (i_t - a_r - a\pi_{t+12} - b\pi^e_{t,12})^2 \to \min$$

Minimizing the criterion gives rise to the set of normal equations:

$$\begin{aligned} 0 &= \sum 2(i_t - a_r - a\pi_{t+12} - b\pi^e_{t,12})(-1) \\ 0 &= \sum 2(i_t - a_r - a\pi_{t+12} - b\pi^e_{t,12})(-\pi_{t+12}) \\ 0 &= \sum 2(i_t - a_r - a\pi_{t+12} - b\pi^e_{t,12})(-\pi^e_{t,12}) \end{aligned}$$

which can be conveniently rewritten as

$$\begin{aligned} 0 &= \bar{i} - \hat{a}\bar{\pi} - \hat{b}\bar{\pi}^e \\ 0 &= \mathrm{cov}(i_t, \pi_{t+12}) - \hat{a}\,\mathrm{var}(\pi_{t+12}) - \hat{b}\,\mathrm{cov}(\pi_{t+12}, \pi^e_{t,12}) \\ 0 &= \mathrm{cov}(i_t, \pi^e_{t+12}) - \hat{a}\,\mathrm{cov}(\pi_{t+12}, \pi^e_{t+12}) - \hat{b}\,\mathrm{var}(\pi^e_{t,12}) \end{aligned}$$

where $\bar{x} = \dfrac{1}{N}\sum_{t=1}^{N} x_t$. Or in a matrix form:

$$\begin{bmatrix} \hat{a} \\ \hat{b} \end{bmatrix} \begin{bmatrix} \mathrm{var}(\pi_{t+12}) & \mathrm{cov}(\pi_{t+12}, \pi^e_{t,12}) \\ \mathrm{cov}(\pi^e_{t,12}, \pi_{t+12}) & \mathrm{var}(\pi^e_{t,12}) \end{bmatrix} = \begin{bmatrix} \mathrm{cov}(i_t, \pi_{t+12}) \\ \mathrm{cov}(i_t, \pi^e_{t,12}) \end{bmatrix}$$

where $\begin{bmatrix} \mathrm{var}(\pi_{t+12}) & \mathrm{cov}(\pi_{t+12}, \pi^e_{t,12}) \\ \mathrm{cov}(\pi^e_{t,12}, \pi_{t+12}) & \mathrm{var}(\pi^e_{t,12}) \end{bmatrix}$ is the information matrix **X'X**.

Solving for the parameter estimates gives

$$\hat{b} = \frac{1}{1-\rho^2}\,\frac{\mathrm{cov}(i_t, \pi^e_{t,12})}{\mathrm{var}(\pi^e_{t,12})} - \frac{\rho}{1-\rho}\,\frac{\mathrm{cov}(i_t, \pi_{t+12})}{\mathrm{std}(\pi_{t+12})\mathrm{std}(\pi^e_{t,12})}$$

$$\hat{a} = \frac{\mathrm{cov}(i_t, \pi_{t+12})}{\mathrm{var}(\pi_{t+12})} - \hat{b}\rho\frac{\mathrm{std}(\pi^e_{t,12})}{\pi_{t+12}}$$

$$\hat{a}_r = \bar{i} - \hat{a}\bar{\pi} - \hat{b}\bar{\pi}^e$$

where $\rho = \dfrac{\mathrm{cov}(\pi_{t+12}, \pi^e_{t,12})}{\mathrm{std}(\pi_{t+12})\mathrm{std}(\pi^e_{t,12})}$[7]

[7] Note that $\rho = 0$ gives rise to a standard OLS estimate.

Next, to evaluate the impact of collinearity on the t-test, we have to analyze its effect on parameters' standard errors. For simplicity, let us assume that regression residuals are homoscedastic and uncorrelated. Then the parameters' variance can be expressed as:

$$\operatorname{var}\begin{bmatrix}\hat{a}\\ \hat{b}\end{bmatrix} = \sigma^2\,(\mathbf{X'X})^{-1} = \sigma^2\begin{bmatrix}\operatorname{var}(\pi_{t+12}) & \operatorname{cov}(\pi_{t+12},\,\pi^e_{t,12})\\ \operatorname{cov}(\pi^e_{t,12},\pi_{t+12}) & \operatorname{var}(\pi^e_{t,12})\end{bmatrix}^{-1} =$$

$$= \sigma^2\begin{bmatrix}\operatorname{var}(\pi_{t+12}) & \rho\,\operatorname{std}(\pi_{t+12})\operatorname{std}(\pi^e_{t,12})\\ \rho\,\operatorname{std}(\pi^e_{t,12})\operatorname{std}(\pi_{t+12}) & \operatorname{var}(\pi^e_{t,12})\end{bmatrix}^{-1} =$$

$$= \frac{\sigma^2}{(1-\rho^2)\,\operatorname{var}(\pi_{t+12})\operatorname{var}(\pi^e_{t,12})}\begin{bmatrix}\operatorname{var}(\pi_{t+12}) & \rho\,\operatorname{std}(\pi_{t+12})\operatorname{std}(\pi^e_{t,12})\\ \rho\,\operatorname{std}(\pi^e_{t,12})\operatorname{std}(\pi_{t+12}) & \operatorname{var}(\pi^e_{t,12})\end{bmatrix}$$

$$\operatorname{var}(\hat{a}_r) = \bar{\pi}^2\operatorname{var}(\hat{a}) + \bar{\pi}^{e2}\operatorname{var}(\hat{b}).$$

The analysis is limited for the above case only. Because the autocorrelation is accounted for in the critical values, this limitation is suitable for further purposes.

To quantitatively analyze the influence of collinearity on the parameters' estimates and t-tests, their values are simulated for different magnitudes of correlation (ρ) between π_{t+12} and $\pi^e_{t,12}$. The analysis is conducted on the set of descriptive statistics summarized in *Table 7*.

TABLE 7 The Calibrated Values – Actual Data Statistics

Statistics	**Estimate**
ρ	0.5957
$\operatorname{cov}(i_t,\ \pi_{t+12})$	1.6795
$\operatorname{cov}(i_t,\ \pi^e_{t,12})$	0.7405
$\operatorname{std}(\pi_{t+12})$	2.0575
$\operatorname{std}(\pi^e_{t,12})$	0.7188
$\operatorname{var}(\pi_{t+12})$	4.2332
$\operatorname{var}(\pi^e_{t,12})$	0.5167
$\bar{i}$	5.2000
$\bar{\pi}$	2.5432
$\bar{\pi}^e$	4.0341

From the graphical results (Figures 5–7) follows that only the parameter a is sensitive to the value of ρ. Its estimate differs considerably for $\rho = 0$ and $\rho = 6$ and so do the t-statistics. However, what is important here is that the parameter is not statistically significant for $\rho = (0.5, 0.9)$. By the nature of expectations, the correlation between $\pi^e_{t,12}$ and π_{t+12} should not be very low, even though the REH does not hold.

Parameter a_r is always statistically insignificant and b is statistically significant under the simulation setup. Hence, it might be concluded the results presented in the paper are robust for $\rho = (0.5, 0.9)$.

FIGURE 5 Dependence of $\hat{a}_r$ and its t-statistics on ρ

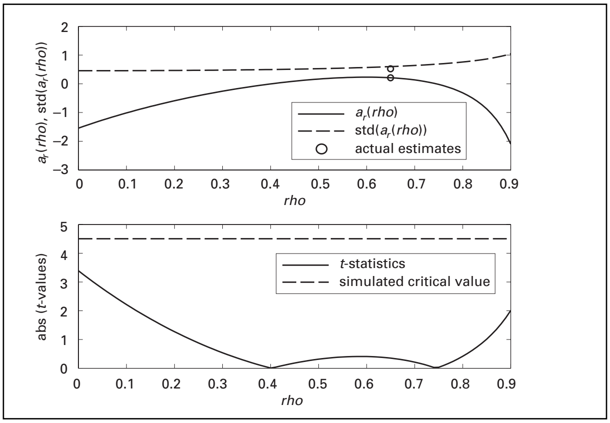


FIGURE 6 Dependence of $\hat{a}$ and its t-statistics on ρ

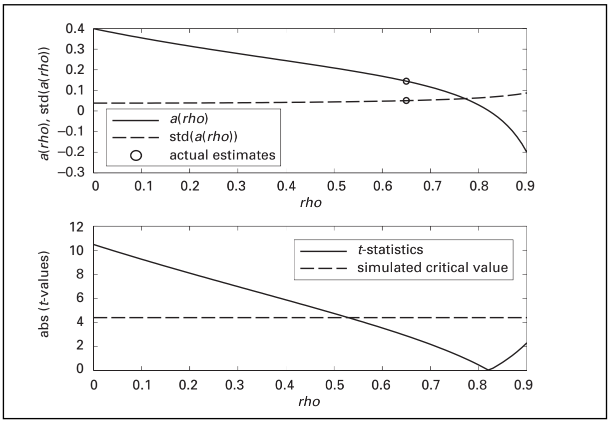
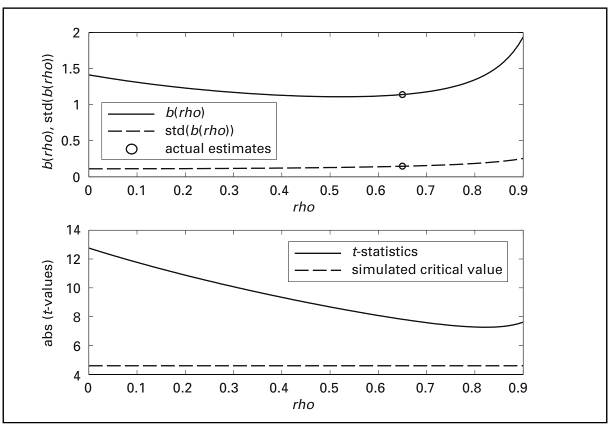


FIGURE 7 Dependence of $\hat{b}$ and its t-statistics on ρ



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SUMMARY

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Do the Measurements of Financial Market Inflation Expectations Yield Relevant Macroeconomic Information?

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Monthly data concerning the inflation expectations of financial analysts in the Czech Republic exhibit a tendency for bias and ineffectiveness. This paper analyses, from a macroeconomic perspective, whether the surveyed data include any relevant macroeconomic information, specifically, whether the surveyed expectations correspond to market expectations considered in macroeconomic analysis and models. Using a methodology based on a simple Fisher rule, it is found that the difference between the surveyed and market inflation expectations is not statistically significant. From this perspective, it is concluded the surveyed inflation expectations bear economically relevant information.