A factor-augemented model of markup on mortgage loans in Poland

Victor Bystrov

University of Lodz

9 September 2013

A Factor-Augmented Model of Markup on Mortgage Loans in Poland

Abstract

The paper describes the results of estimation of a factor-augmented vector autoregressive model that relates the markup on mortgage loans in national currency, granted to households by monetary financial institutions, and 1-month inter-bank rate that represents the cost of funds for financial institutions. The factors by which the model is augmented, summarize information that can be used by banks to forecast interest rates and evaluate macroeconomic risks. The estimation results indicate that there is a significant relation between the markup and the changes in 1-month WIBOR. This relation can be interpreted as evidence of incomplete transmission of the monetary policy shocks to mortgage rates set by monetary financial institutions. The policy shocks are partially absorbed by changes in the markup.

JEL: C32, C53, E43, E44

Keywords: factor models, interest rates, pass-through, markup

1 Introduction

An interest rate pass-through model represents a relation between retail rates set by monetary financial institutions for households and firms and a policy rate or a wholesale market rate representing cost of funds for financial institutions. Before the emergence of the financial crisis, the empirical literature on the interest rate pass-through focused on the estimation of bivariate models relating a retail rate (e.g., mortgage rate) and a market rate (e.g., EURIBOR). The extent and the speed of the pass-through were considered as indicators of the effectiveness of monetary policy.

Under conditions of financial turmoil, the effect of policy rates and short-term market rates on retail rates has become weaker, and the conventional models of the pass-through have become poor representations of the monetary policy transmission. An augmentation of the conventional models is needed in order to account for additional information that monetary financial institutions consider when settting retail rates for households and firms.

In this paper we consider a factor-augmented vector autoregressive (FAVAR) model that explains deviations from the long-run equilibrium defined by a conventional model of the pass-through for mortgage rates in Poland. A conventional model is based on the assumption of a constant markup of a retail rate over a wholesale rate. Persistence changes in the markup imply deviations from the equilibrium. The FAVAR estimated in this paper, measures relations between the markup, changes in a wholesale rate, and a few common factors that are estimated using a large panel of macroeconomic and financial indicators.

The empirical model is motivated by a theoretical forward-looking model which describes a relation between the markup and the expectations formed by monetary financial institutions. The common factors summarize information that can be used by monetary financial institutions in the evaluation of macroeconomic risks and the forecasting of future interest rates.

The paper is organized as follows. In the next section we describe a simple theoretical model of the pass-through. Section 3 includes a description of the econometric model. Data description is given in Section 4. The estimation results are reported in Section 5. Section 6 concludes the paper.

2 Aggregation, expectations and the pass-through

Monetary Financial Institutions (MFI) Interest Rates (MIR) statistics adopted in the EU countries, including Poland, provides synthetic retail bank rates that are aggregated into a few broad categories defined by the type of a product and its maturity (e.g., loans for house purchases over 1 year and up to 5 years maturity). The aggregation is performed by reporting agents (monetary financial institutions). Therefore, no systematic statistical data are available for individual products of exact maturity.

The economic literature on the interest rate pass-through uses these retail rates to match them with money market rates or government bond yields (wholesale rates) defined for specific maturities (see, e.g., de Bondt 2005). As there is no exact matching of maturities between retail rates and wholesale rates, two approaches are commonly used: either a retail rate is matched to a short-term money market rate approximating a policy rate (like 1 or 3-month EURIBOR), or an appropriate wholesale rate is chosen on the basis of correlation analysis among those rates which are closest to a given retail rate in maturity. A notable exception is the study by Sorensen and Werner (2006) who construct synthetic wholesale rates.

The first approach ignores the maturity transformation and is only valid if there is a stable relation between short-term money market rates and long-term bond yields. The second approach uses an ad hoc method which may match different wholesale rates over different sub-samples of data.

A MIR rate is a synthetic rate representing a weighted average of retail rates of various maturities:

$$r_t = \sum_{\tau=\underline{\tau}}^{\overline{\tau}} \omega_\tau r_t(\tau),$$

where $r_t(\tau)$ is a retail rate of maturity τ (here, maturity means period of interest rate fixation), $\underline{\tau}$ is the minimal maturity and $\overline{\tau}$ is the maximal maturity of retail rates which are included in the synthetic rate r_t, ω_τ is the weight of a rate of maturity τ. The weights ω_τ, $\tau = \underline{\tau}, \underline{\tau} + 1, ...\overline{\tau}$, are not systematically reported by MFIs.

If monetary financial institutions were matching maturities of retail and wholesale rates, then the baseline pass-through equation would have the form,

$$r_t = \nu + \beta \sum_{\tau=\underline{\tau}}^{\overline{\tau}} \omega_\tau m_t(\tau), \tag{1}$$

where $m_t(\tau)$ is a wholesale rate on a debt obligation of maturity τ, β is the pass-through coefficient, and ν is the bank mark-up. The parameters ν and β are said to be determined by the demand elasticity and the market structure (de Bondt 2005). If $\beta < 1$, then the pass-through is said to be incomplete. The incomplete pass-through is explained by microeconomic factors such as low market competitiveness and credit rationing (see, inter alia, Winker 1999; Kot 2003; Chmielewski 2004; Gambacorta 2006; Sorensen and Werner 2006). However, in this paper we consider a macroeconomic model of the incomplete pass-through.

If MFIs do not match maturities, but rely on the short-term financing, then they have to forecast a short-term wholesale rate and determine a risk premium in order to set a retail rate. Let us consider a modification of the linearized expectations model proposed by Shiller (1979), which relates a wholesale rate $m_t(\tau)$ of maturity τ to the expected path of a one-period rate, $m_t(1)$:

$$m_t(\tau) = \phi_t(\tau) + \frac{1-\gamma}{1-\gamma^\tau} \sum_{h=0}^{\tau-1} \gamma^h E_t m_{t+h}(1), \tag{2}$$

where $\phi_t(\tau)$ is a time-varying risk premium, γ is a discount factor, and $E_t m_{t+h}(1)$ is the expectation of the one-period rate.

Substituting (2) in (1), we obtain

$$r_t = \nu + \beta \sum_{\tau=\underline{\tau}}^{\overline{\tau}} \omega_\tau \left[\phi_t(\tau) + \frac{1-\gamma}{1-\gamma^\tau} \sum_{h=0}^{\tau-1} \gamma^h E_t m_{t+h}(1) \right].$$

After rearrangement, using $\Delta^{(h)} m_{t+h}(1) = m_{t+h}(1) - m_t(1)$:

$$r_t = \left[\nu + \beta \sum_{\tau=\underline{\tau}}^{\overline{\tau}} \omega_\tau \phi_t(\tau) \right] + \beta m_t(1) + \beta \left[\sum_{\tau=\underline{\tau}}^{\overline{\tau}} \omega_\tau \frac{1-\gamma}{1-\gamma^\tau} \sum_{h=1}^{\tau-1} \gamma^h E_t \Delta^{(h)} m_{t+h}(1) \right].$$

The equation can be rewritten as

$$r_t = \mu_t + \beta m_t(1) + \sum_{h=1}^{\overline{\tau}-1} \delta_h E_t \Delta^{(h)} m_{t+h}(1),$$

where

$$\mu_t = \left[\nu + \beta \sum_{\tau=\underline{\tau}}^{\overline{\tau}} \omega_\tau \phi_t(\tau) \right], \text{ and } \delta_h = \begin{cases} \beta \sum_{\tau=\underline{\tau}}^{\overline{\tau}} \omega_\tau \frac{1-\gamma}{1-\gamma^\tau} \gamma^h, & h \leq \underline{\tau} - 1 \\ \beta \sum_{\tau=h+1}^{\overline{\tau}} \omega_\tau \frac{1-\gamma}{1-\gamma^\tau} \gamma^h, & h \geq \underline{\tau} \end{cases}.$$

The synthetic retail rate r_t can be expressed as a function of a spot short-term rate and expected changes in the short term rate up to the maximal maturity of retail products included in the synthetic rate r_t. The residual variability of the retail rate can be explained by the fluctuations in the risk premium.

In the Polish market of mortgage loans, the predominant pricing mechanism for loans granted in national currency is to set a retail rate equal to a short-term

WIBOR (Warsaw Inter-Bank Offered Rate) plus a markup. Such pricing mechanism implies that the long-term value of the pass-through coefficient β should be equal to one. The markup, defined as a difference between the retail rate on mortgage loans and a 1-period WIBOR, is given by

$$z_t = r_t - m_t(1) = \mu_t + \sum_{h=1}^{\overline{\tau}-1} \delta_h E_t \Delta^{(h)} m_{t+h}(1),$$

where $\overline{\tau}$ is the maximal period of interest rate fixation.

In this model, persistent changes in the markup z_t are caused by changes in the evaluation of risk and revisions of forecasts by monetary financial institutions. The persistent changes in the markup mean ineffectiveness of the monetary policy based on the regulation of interest rates, as monetary financial institutions do not fully transmit changes in wholesale (market) rates to retail rates, but partially absorb those changes through changes in the markup.

3 Econometric Model

In this paper a dynamic factor model is employed to summarize information that can be used by MFIs in making projections of future interest rates and evaluation of risk. In a similar study, Banerjee, Bystrov and Mizen (2013) estimated a pass-through model where recursive forecasts of a market rate were included into a dynamic regression. The forecasts were based on a factor model of the yield curve. Though the study confirmed the importance of forecasts in the retail rate setting, the forecasts were based on the information contained in the yield curve only and the risk premium was assumed to be constant. The performance of the model may be improved by extending the information set with other macroeconomic and financial variables that might be useful in the forecasting of market rates and the evaluation of risk.

We model expectations of MFIs as based on a dynamic factor model which represents an extensive set of macroeconomic and financial variables by a few common factors:

$$X_t = \Lambda_t F_t + e_t,$$

where X_t is $(N \times 1)$ vector of observed stationary macroeconomic and financial indicators, F_t is $(R \times 1)$ vector of unobserved common factors $(R << N)$, Λ_t is $(N \times R)$ matrix of loadings, and e_t is $(N \times 1)$ vector of idiosyncratic components.

A factor forecast can be constructed as a direct projection on the estimated common factors, their lags and lags of the forecast variable:

$$\Delta^{(h)}\widehat{m}_{t+h|t}(1) = a_t^{(h)} + \sum_{k=0}^{K} b_{kt}^{(h)\prime} \widehat{F}_{t-k}^{(t)} + \sum_{l=0}^{L} c_{lt}^{(h)} \Delta m_{t-l}(1), \;\; h = 1, 2, ..., \tau - 1,$$

where $\Delta^{(h)}\widehat{m}_{t+h|t}(1)$ is a forecast of h-period change in a one-period market rate conditional on the information available at time t.

MFIs can use a variety of macroeconomic and financial indicators to forecast future interest rates. The information contained in these indicators can be parsimoniously summarized by few common factors, and the factor forecasts can serve an approximation of the expectations formed by MFIs. However, the information, summarized by the common factors, may also be used in the evaluation of macroeconomic risks and the determination of the risk premium (μ_t).

Therefore, the inclusion of factor forecasts in a pass-through model, while assuming a constant risk premium, may not be recommended. The risk premium can be time-varying and dependent on the macroeconomic indicators which are included in the factor model. In this paper, a dynamic model of the pass-through is augmented by the estimated common factors, assuming that the information, which is summarized by these factors, may determine both expectations and the risk premium.

A bivariate vector autoregression, including a measure of markup, z_t, and a change in a market rate, $\Delta m_t(1)$, is augmented by a few factors $\widehat{F}_t$ extracted from a large number of macroeconomic and financial indicators:

$$\begin{bmatrix} \widehat{F}_t^{(T)} \\ \Delta m_t(1) \\ z_t \end{bmatrix} = \begin{bmatrix} \Phi(L) & 0 & 0 \\ \alpha(L)' & a(L) & b(L) \\ \beta(L)' & c(L) & d(L) \end{bmatrix} \begin{bmatrix} \widehat{F}_{t-1}^{(T)} \\ \Delta m_{t-1}(1) \\ z_{t-1} \end{bmatrix} + \begin{bmatrix} \eta_t \\ \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix},$$

where $\widehat{F}_t^{(T)}$ is $(R \times 1)$ vector of estimated common factors; $\Phi(L)$ is $(R \times R)$ matrix lag polynomial; $\alpha(L)$ and $\beta(L)$ are $(R \times 1)$ vector polynomials; $a(L)$, $b(L)$, $c(L)$, and $d(L)$ are scalar polynomials. The common factors F_t are assumed to be exogenous with respect to markup z_t and differenced market rate $\Delta m_t(1)$. Therefore, zero restrictions are imposed on the lags of z_t and $\Delta m_t(1)$ in the equations for the common factors.

The common factors are estimated using the principal components estimator. Bai and Ng (2006) provide central limit theorems and confidence intervals for inference in factor-augmented regressions. We implement Bai and Ng (2006) methodology: parameters of the factor-augmented regressions are estimated using the least squares estimator and heteroscedasticity-consistent standard errors are computed to account for consequences of including generated regressors in the model.

4 Data Description

The FAVAR model is estimated using monthly data from January 2004 to December 2012. The markup is computed as a difference between the average rate on outstanding amounts on mortgage loans granted in national currency and the monthly average of 1-month WIBOR. It is a synthetic measure of markup that can only be interpreted as an approximation of the actual markup set by MFIs. The mortgage rate is extracted from the Monetary and Financial Statistics of the National Bank of Poland. Figure 1 shows the time series plot of two series (with means subtracted).

The dynamic factor model is estimated using monthly data from January 2001 to December 2012. The data include 49 time series covering industrial production, prices, exchange rates, interest rates, monetary aggregates, stock exchange indices and leading business indicators (see Table 1 in the Appendix). The series were extracted from a databases of a few institutions: the National Bank of Poland (NBP), the Central Statistical Office of Poland (GUS), the Warsaw Stock Exchange (GPW), Eurostat, the European Central Bank (ECB), the OECD, and the IMF. The composition of the data panel is aimed to provide a balanced representation of all sectors of the Polish economy.

Prior to estimation of the factor model, the data were processed using Stock and Watson (1999) methodology. First, all series that were modelled as generated by integrated processes, were transformed to stationary series, using differences or log-differences. Second, all non-financial time series were seasonally adjusted using Census X12-ARIMA procedure. Third, outliers exceeding the interquartile difference by a factor of six, were substituted by missing values that were subsequently substituted by estimates obtained using the expectation-maximization (EM) algorithm. Fourth, all series were standardized to have zero mean and unit variance.

5 Estimation Results

The common factors were estimated using the principal component estimator. First, the estimation was performed a panel of series including no missing values. Second, missing values were interpolated using the EM algorithm and the factors were estimated using the whole panel.

Initially, ten common factors were estimated. Of those, six factors were selected using a threshold of at least 5 percent of the total variance explained by each selected factor (for an application of such criterion, see Forni and Reichlin 1998). The final FAVAR obtained in the model selection process, included only four factors. These four factors explain 50 percent of the total variance in 49 time series. The time series plot of these factors is shown in Figure 2 and the loadings of these factors onto individual time series are presented in Figures 3-4. The first factor loads on nominal indicators: producer price inflation, returns on exchange rates and stock indices. The second factor loads on indicators of industrial production and foreign trade. The third factor is correlated with interest rates and the fourth factor has the highest correlation with indicators of consumer price inflation.

The selection of the FAVAR model was carried out using System SER (Sequential Elimination of Regressors) procedure based on the Bayesian Information Criterion (Brüggemann and Lütkepohl 2001). The initial model included six factors and six lags of factors, the markup and the difference of 1-month WIBOR. The final model included four factors and up to three lags of each variable (see

Table 2).

The eigenvalues of the companion matrix of the estimated FAVAR and the covariance matrix of disturbances are reported in Table 3. All eigenvalues of the companion matrix are less than one in modulus, which means that the estimated FAVAR is stable (there are no unit or explosive roots).

The orthogonalized impulse responses of the differenced 1-month WIBOR and the mortgage markup are presented in Figures 5 and 6. The figures show the 95 percent joint bootstrap confidence bands obtained using the neighbouring paths method proposed by Staszewska (2007) (see also Staszewska-Bystrova 2011). The joint band contains the entire response function constructed for a given response horizon with probability 0.95.

The assumed ordering of variables in the FAVAR means that common factors can have an immediate effect onto the 1-month WIBOR and the markup. However, the 1-month WIBOR and the markup have no immediate effect on the common factors. This is consistent with the interpretation of the common factors as exogenous latent variables that describe a state of the economy. The ordering of the variables also implies that the 1-month WIBOR may have an immediate impact onto the markup, but not vice versa.

Figure 5 shows the orthogonalized impulse responses of the markup, z_t, to shocks in the common factors F_{1t}, F_{2t}, F_{3t}, and F_{4t}, and in the differenced 1-month WIBOR, $\Delta m_t(1)$. Each shock is equal to one standard deviation of residuals in the estimated equation for a corresponding variable. An impulse to the differenced 1-month WIBOR, $\Delta m_t(1)$, implies a significant negative shock to the markup, z_t, which slowly converges to the previous level afterwards. It means that if the level of 1-month WIBOR unexpectedly but permanently increases, then the markup decreases temporarily: a part of the increase in the WIBOR is not transferred to an increase in mortgage rates - it is absorbed by a decrease in the markup.

There are significant impulse responses of the markup to factors 2, 3, and 4, which are correlated with the growth rates of industrial production, changes in interest rates, and consumer price inflation correspondingly. The direction of these effects should be interpreted with a precaution though, as factors are identified up to a linear transformation, and additional restrictions have to be imposed to admit a structural interpretation of these impulse responses. However, it can be concluded that one has to control for other macroeconomic indicators when measuring the response of the markup to a short-term market rate.

The orthogonal impulse responses of the 1-month WIBOR are shown in Figure 6. There are significant responses of the differenced 1-month WIBOR to shocks in factors correlated with growth rates of industrial production, changes in other interest rates, and consumer price inflation. These factors summarize information of potential use in the forecasting of 1-month WIBOR and the determination of the markup which is set on mortgage loans by MFIs. There is no feedback from the markup to 1-month WIBOR, as lags of the markup are excluded from the equation for the differenced WIBOR by the model selection procedure, and the markup is assumed to have no instantaneous effect on the WIBOR.

In order to investigate the parameter stability of the estimated FAVAR model, we implement the Andrews (1993) and the Andrews-Ploberger (1994) tests, which are based on the supremum Wald statistic and the average exponential Wald statistic respectively. It was demonstrated in several studies (see Stock and Watson 1998; Hansen 2000; Cogley and Sargent 2005) that these tests have the highest power, as compared to other popular tests, to detect parameter instability in dynamic regressions.

Following Stock and Watson (1996), we implement a heteroscedasticity-robust version of these tests and compute bootstrapped critical values to account for a small sample size. The Andrews the Andrews-Ploberger tests are not based on the assumption of a specific date of structural change, but evaluate stability of the parameters over a window of observations. We selected a symmetric window, trimming 25 percent of the observations at the beginning and at the end of the estimation sample. As a result, the stability of parameters was tested over the period from May 2006 to September 2010. A choice of a larger window would lead to computational instability of the least square estimator.

Table 4 reports computed test statistics together with bootstrapped and asymptotic critical values for the 5 percent level of significance. The results are provided for each equation of the FAVAR model and for the model as a whole. The test statistics of both tests are smaller than bootstrapped and asymptotic critical values for all equations of the model. Therefore, the null hypothesis of parameter stability cannot be rejected given the significance level of 5 percent. No evidence of parameter instability in the FAVAR model is found.

6 Conclusions

In this paper, a factor-augmented VAR model is used to explain the relation between changes in 1-month WIBOR and the markup on mortgage loans in national currency, which are granted to households by monetary financial institutions. The augmentation of a simple VAR model is motivated by the forward-looking behaviour of monetary financial institutions that determine the markup on mortgage loans in dependence on their projections of future interest rates and their evaluation of risk. The common factors, which are computed using a large panel of economic and financial time series, summarize, in a parsimonious way, the information used by monetary financial institutions.

The estimation results confirm that there is a significant relation between the markup and changes in 1-month WIBOR. It implies that monetary policy shocks transmitted to changes in 1-month WIBOR are only partially transmitted to changes in rates paid by households on housing loans. The changes in 1-month WIBOR are partially absorbed by changes in the markup. The common factors are found to be significant in the model, which means that additional information summarized by the factors influences the pricing behavior of financial institutions.

A few extensions of the research are possible. First, additional assumptions can be imposed and a structural FAVAR can be estimated in order to provide a more profound economic interpretation of the impulse response analysis. Second, instead of considering heteroscedasticity-robust estimator, the direct modelling of heteroscedasticity can be implemented. Third, an explicit measure of the risk premium can be derived, using the dynamic factor model, and its effect onto the markup can be evaluated.

7 Literature

Andrews D. W. (1993), Testing for parameter instability and structural change with unknown change point, Econometrica, Vol. 61, No. 4, pp. 821-856.

Andrews D. W., Ploberger W. (1994), Optimal tests when a nuisance parameter is presented only under the alternative, Econometrica, Vol. 62, No. 6, pp. 1383-1414.

Bai J., Ng S. (2006), Confidence intervals for diffusion index forecasts and inference for factor-augmented regressions, Econometrica, Vol. 74, No. 4, pp. 1133-1150.

Banerjee A., Bystrov V., Mizen P. (2012), How do anticipated changes to short-term market rates influence banks' retail rates? Evidence from the four major euro area economies, Forthcoming in Journal of Money, Banking and Credit.

Brüggemann R., Lütkepohl, H. (2001), Lag selection in subset VAR models with an application to a U.S. monetary system, in R.Friedmann, L. Küppel and H. Lütkepohl (eds.), Econometric Studies: a Festscgrift in Honour of Joachim Frohn, LIT Verlag, Müster, pp. 107-128.

Chmielewski T. (2004), Interest rate pass-through in the Polish banking sector and bank-specific financial disturbances, In: Papers Presented at the ECB Workshop on Asset Prices and Monetary Policy, Frankfurt, December 11-12, 2003.

de Bondt G. (2005), Interest rate pass through in the euro area, German Economic Review, 6, pp. 37-78.

Forni M., Reichlin L. (1998), Let's Get Real: A Factor Analytical Approach to Disaggregated Business Cycle Dynamics, Review of Economic Studies, 65, 3, pp. 453-473.

Gambacorta L. (2008), How Do Banks Set Interest Rates?, European Economic Review, 52, pp 792-819.

Hansen B. E. (2000), Testing for structural change in conditional models, Journal of Econometrics, 97, 1, pp. 93-115.

Kok-Sørensen, C., Werner T. (2006), Bank interest rate pass through in the euro area, ECB Working Paper No 580.

Kot A. (2004), Is interest rates pass-through related to banking sector competitiveness? Paper presented at the Third Macroeconomic Policy Research Workshop on Monetary Transmission in the New and Old Members of the EU, October 29-30, 2004, Budapest.

Cogley T., Sargent T. J. (2005), Drifts and volatilities: monetary policy and outcomes in the post WWII US, Review of Economic Dynamics, 8, pp. 262-302.

Shiller R. J. (1979), The volatility of long-term interest rates and the expectations models of the term structure, Journal of Political Economy, Vol. 87, No. 6, pp. 1190-1219.

Staszewska A. (2007), Representing uncertainty about response paths: the use of heuristic optimization methods, Computational Statistics and Data Analysis, 52, 1, pp. 121-132.

Staszewska-Bystrova A. (2011), Bootstrap prediction bands for forecast paths from vector autoregressive models, Journal of Forecasting, 30, 8, pp. 721-735.

Stock, J. H., Watson M. W.(1998), Median unbiased estimation of coefficient variance in a time-varying parameter model, Journal of American Statistical Association, Vol. 93, No. 441, pp. 349-358.

Winker P. (1999), Sluggish adjustment of interest rates and credit rationing: an application of unit root testing and error correction modelling, Applied Economics 31, 3, pp. 267-277.

Appendix

Figure 1: Time series plot of markup and Δm_t

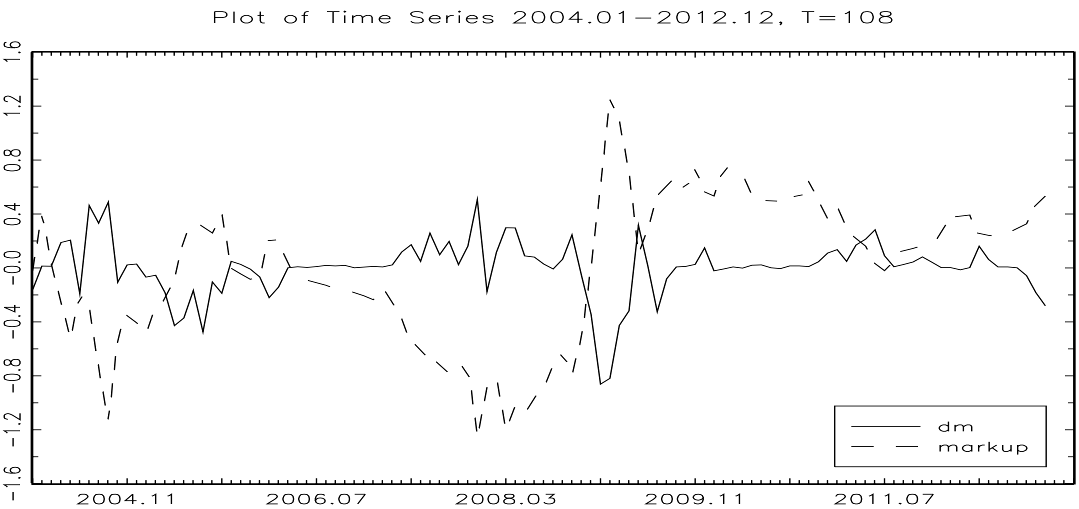


Figure 2: Time series plot of four factors

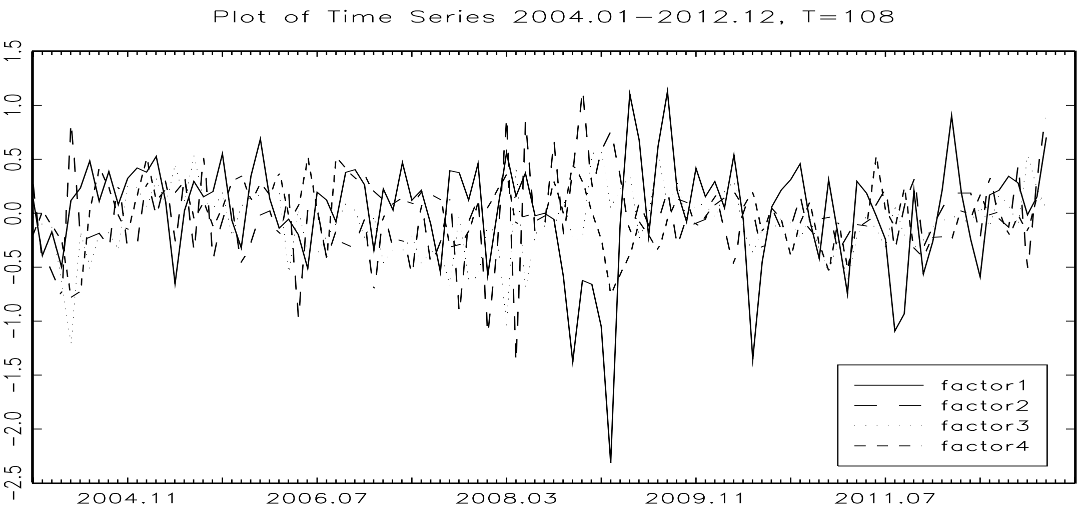


Table 1: Description of data panel

N	Mnemonic	Description	Data Source	SA*	TC**
1	Retail.Vol	Retail trade, index of deflated turnover	Eurostat	Yes	2
2	IP.Manuf	Manufacturing, volume index of production	Eurostat	Yes	2
3	IP.Mining	Mining and quarrying, volume index of production	Eurostat	Yes	2
4	IP.Constr	Construction, volume index of production	Eurostat	Yes	2
5	IP.Non.Dur	Non-durable consumption goods, volume index of production	Eurostat	Yes	2
6	IP.Dur	Durable consumption goods, volume index of production	Eurostat	Yes	2
7	IP.Interm	Intermediate goods, volume index of production	Eurostat	Yes	2
8	IP.Cap	Capital goods, volume index of production	Eurostat	Yes	2
9	IP.Energy	Energy, volume index of production	Eurostat	Yes	2
10	Empl.Manuf	Employment (number of people employed), manufacturing	Eurostat	Yes	2
11	Wages.Manuf	Gross wages and salaries	Eurostat	Yes	2
12	Unempl	Registered unemployment rate	GUS	Yes	1
13	Exports	Total exports, current prices, PLN mln	GUS	Yes	2
14	Imports	Total imports, current prices, PLN mln	GUS	Yes	2
15	PPI.Non.Dur	Producer price index, non-durable consumption goods	Eurostat	Yes	2
16	PPI.Dur	Producer price index, durable consumption goods	Eurostat	Yes	2
17	PPI.Interm	Producer price index, intermediate goods	Eurostat	Yes	2
18	PPI.Cap	Producer price index, capital goods	Eurostat	Yes	2
19	PPI.Manuf	Producer price index, manufacturing	Eurostat	Yes	2
20	PPI.Energy	Producer price index, energy	Eurostat	Yes	2
21	PPI.Mining	Producer price index, mining and quarrying	Eurostat	Yes	2
22	PPI.Constr	Producer price index, construction	Eurostat	Yes	2
23	PPI.Food	Producer price index, food	Eurostat	Yes	2
24	CPI.Food	Consumer price index, food	OECD	Yes	2
25	CPI.All	Consumer price index, all items, non-food, non-energy	OECD	Yes	2
26	HICP.All	Harmonized index of consumer prices, all items	OECD	Yes	2
27	CPI.Energy	Consumer price index, Energy	OECD	Yes	2
28	BCI.Manuf	Indicator of general business tendency climate, manufacturing	GUS***	Yes	1
29	BCI.Constr	Indicator of general business tendency climate, construction	GUS	Yes	1
30	BCI.Trade	Indicator of general business tendency climate, trade	GUS	Yes	1
31	PLNUSD	Exchange rate, PLN/USD, monthly average	NBP	No	2
32	PLNUSD	Exchange rate, PLN/EUR, monthly average	NBP	No	2
33	NEER41	Nominal effective exchange rate - 41 trading partners	Eurostat	Yes	2
34	REER41	Real effective exchange rate - 41 trading partners	Eurostat	Yes	2
35	M1	Monetary aggregate M1, PLN mln	NBP	Yes	2
36	M2	Monetary aggregate M2, PLN mln	NBP	Yes	2
37	M3	Monetary aggregate M3, PLN mln	NBP	Yes	2
38	Mortg.Loans	Outstanding amounts of mortgage loans granted in national currency, PLN mln	NBP	Yes	2
39	WIBOR.ON	Overnight WIBOR, monthly average	Eurostat	No	1
40	WIBOR.1W	1-week WIBOR, monthly average	Eurostat	No	1
41	WIBOR.3M	3-month WIBOR, monthly average	Eurostat	No	1
42	WIBOR.6M	6-month WIBOR, monthly average	Eurostat	No	1
43	WIBOR.12M	Overnight WIBOR, monthly average	Eurostat	No	1
44	Bond.yield.10Y	Government bond yield, 10 years maturity	Eurostat	No	1
45	WIG	Warsaw Stock Exchange Index, monthly average	GPW	No	5
46	WIGBanks	Warsaw Stock Exchange Index, Banks, monthly average	GPW	No	5
47	SP500	Standard and Poors' 500 Index, monthly average	ECB	No	5
48	EuroStoxx50	Dow Jones EuroStoxx 50 Index, monthly average	ECB	No	5
49	Brent.Price	Brent light blend U.K.	IMF	No	5

*Seasonal adjustment: Yes - series was adjusted, No - series was not adjusted
**Transformation code: 0 - no transformation, 1- difference, 2 - log-difference
***Data Sources: GUS - Main Statistical Office of Poland, GPW - Warsaw Stock Exchange, NBP - National Bank of Poland, ECB - European Central Bank, IMF - International Monetary Fund

Figure 3: Loadings of factors 1 and 2 on individual time series

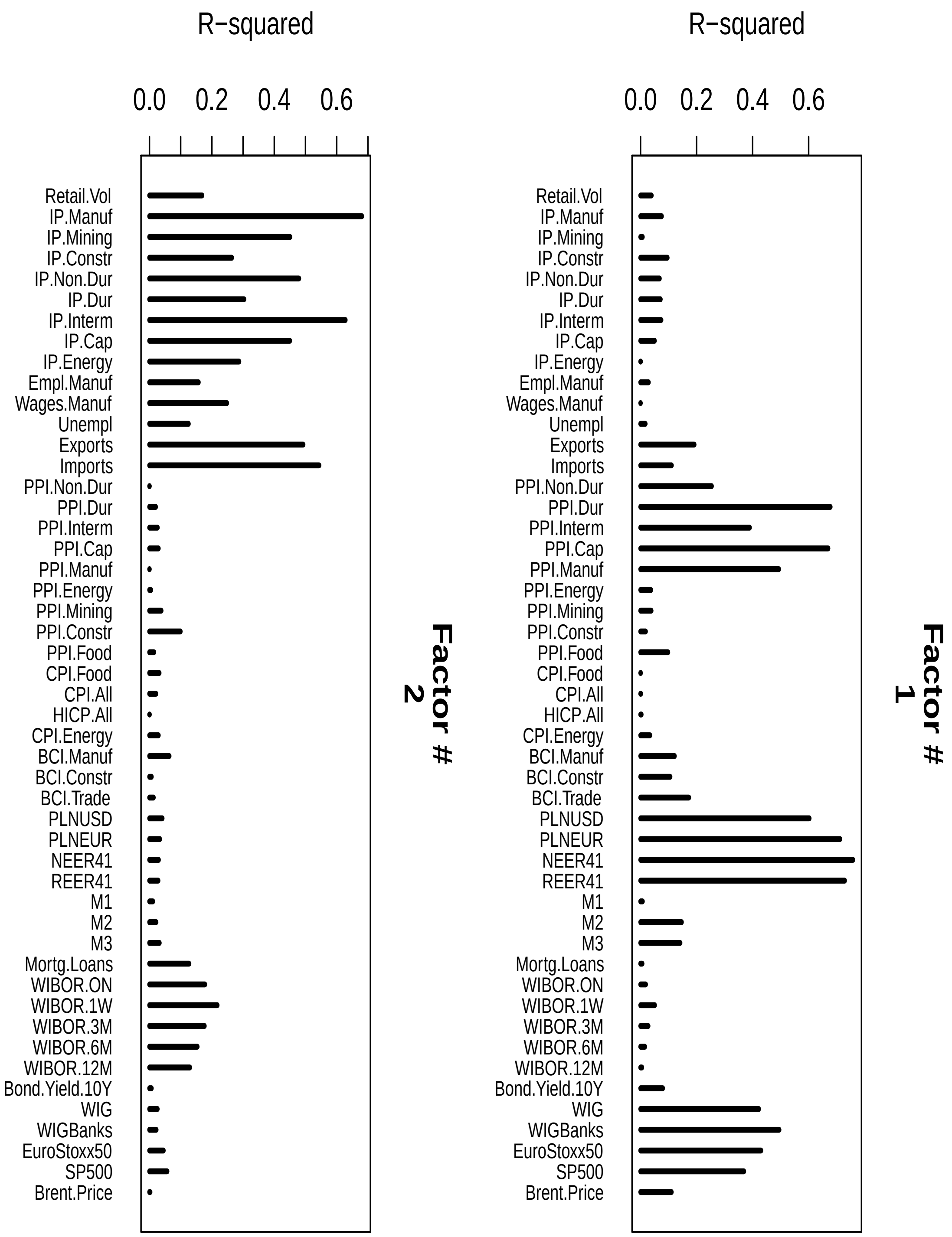
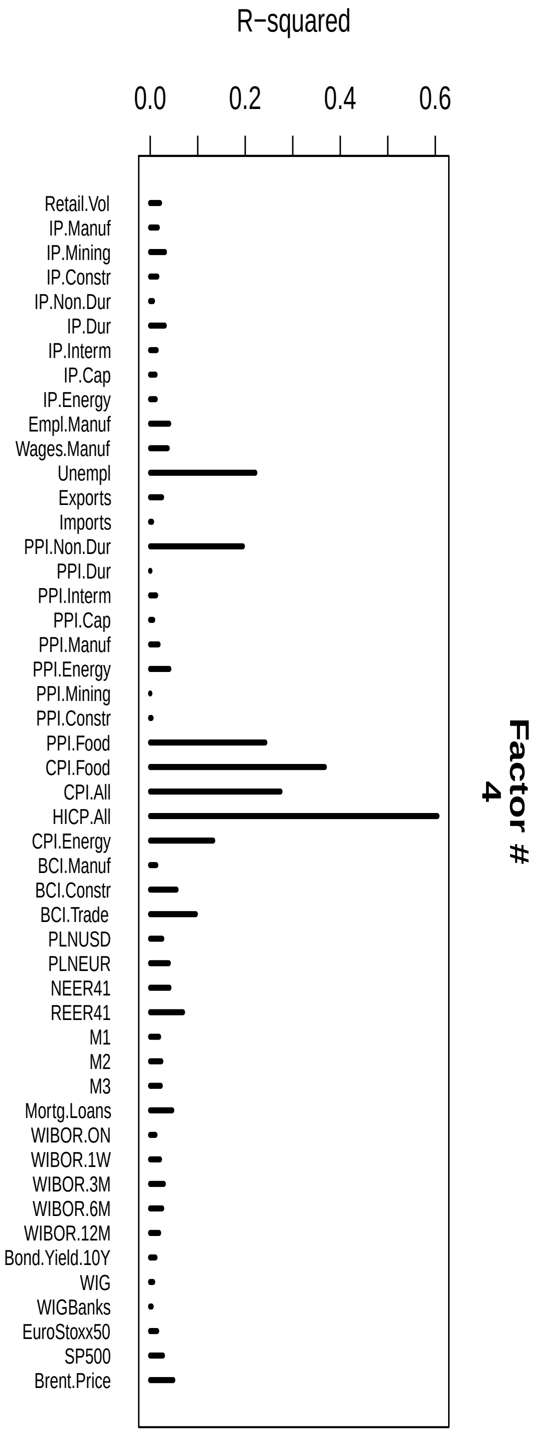


Figure 4: Loadings of factors 3 and 4 on individual time series



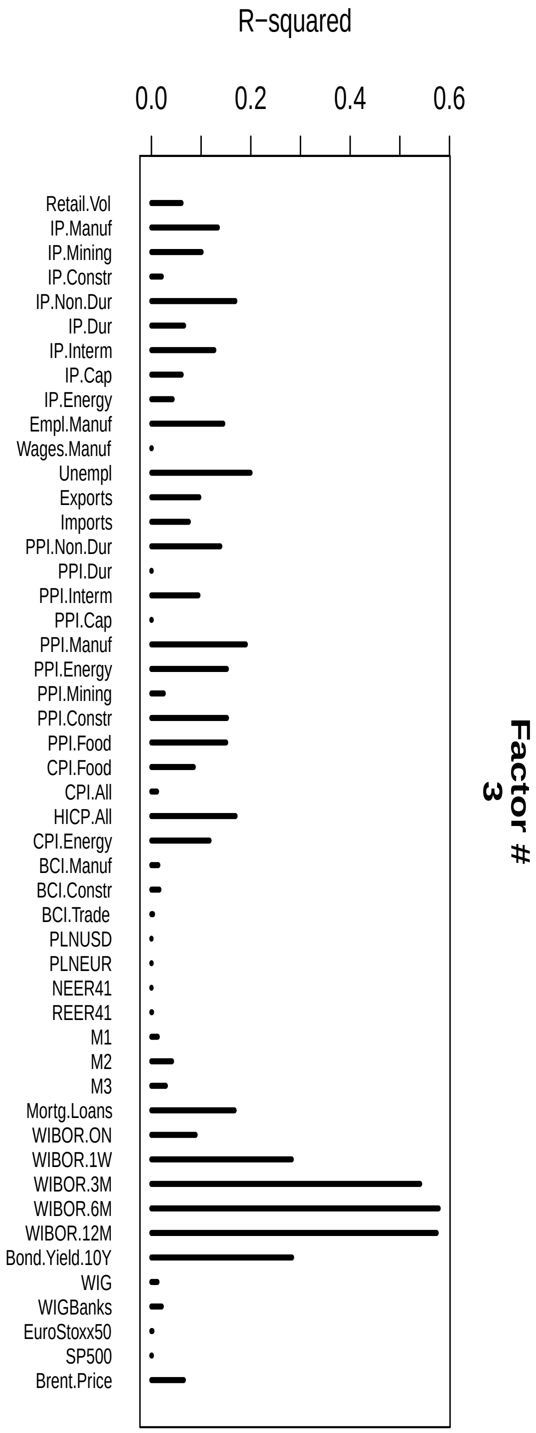


Table 2: Estimates of Factor-Augmented Vector Autoregression

	F_{1t}	F_{2t}	F_{3t}	F_{4t}	Δm_t	z_t
F_{1t-1}	0.328	-0.288	0.105	0.111	0.079	-0.095
	(0.116)*	(0.071)	(0.043)	(0.046)	(0.027)	(0.033)
F_{2t-1}	-0.289	-	0.545	-0.131	-0.223	0.360
	(0.116)	-	(0.064)	(0.074)	(0.033)	(0.048)
F_{3t-1}	0.489	0.842	0.351	0.171	-0.373	0.467
	(0.164)	(0.145)	(0.085)	(0.098)	(0.063)	(0.083)
F_{4t-1}	-	-	-	0.427	0.084	-0.146
	-	-	-	(0.073)	(0.064)	(0.061)
Δm_{t-1}	-	-	-	-	-	-
	-	-	-	-	-	-
z_{t-1}	-	-	-	-	-	0.519
	-	-	-	-	-	(0.088)
F_{1t-2}	-0.306	-	-	-	-	-
	(0.095)	-	-	-	-	-
F_{2t-2}	-0.508	-0.169	0.222	-	-	-
	(0.132)	(0.108)	(0.083)	-	-	-
F_{3t-2}	-	-	-	-	-	-
	-	-	-	-	-	-
F_{4t-2}	-0.520	-	-	-	-	-
	(0.163)	-	-	-	-	-
Δm_{t-2}	-	-	-	-	-	-
	-	-	-	-	-	-
z_{t-2}	-	-	-	-	-	-
	-	-	-	-	-	-
F_{1t-3}	-	-	-	-	-	-
	-	-	-	-	-	-
F_{2t-3}	-0.398	-	-	-	-	-
	(0.136)	-	-	-	-	-
F_{3t-3}	-	-0.349	-	-0.140	-	-
	-	(0.105)	-	(0.087)	-	-
F_{4t-3}	-	-	-	-	-0.207	0.191
	-	-	-	-	(0.058)	(0.062)
Δm_{t-3}	-	-	-	-	-	0.408
	-	-	-	-	-	(0.102)
z_{t-3}	-	-	-	-	-	0.389
	-	-	-	-	-	(0.081)
intercept	-	-	-	-	-0.060	0.238
	-	-	-	-	(0.016)	(0.080)

*Heteroscedasticity - consistent standard errors are reported in parenthesis

Table 3: Estimates of Factor-Augmented Vector Autoregression

Estimated Covariance Matrix of Residuals

	F_{1t}	F_{2t}	F_{3t}	F_{4t}	Δm_t	z_t
F_{1t}	0.172	0.010	0.027	-0.001	0.004	-0.007
F_{2t}	0.010	0.101	-0.031	-0.004	-0.004	-0.002
F_{3t}	0.027	-0.031	0.046	0.011	-0.010	0.011
F_{4t}	-0.001	-0.004	0.011	0.047	0.007	-0.007
Δm_t	0.004	-0.004	-0.010	0.007	0.020	-0.016
z_t	-0.007	-0.002	0.011	-0.007	-0.016	0.022

Six Largest Roots of Companion Matrix

0.951	0.788	0.788	0.693	0.693	0.640

Figure 5: Orthogonalized Impulse Responses of Markup

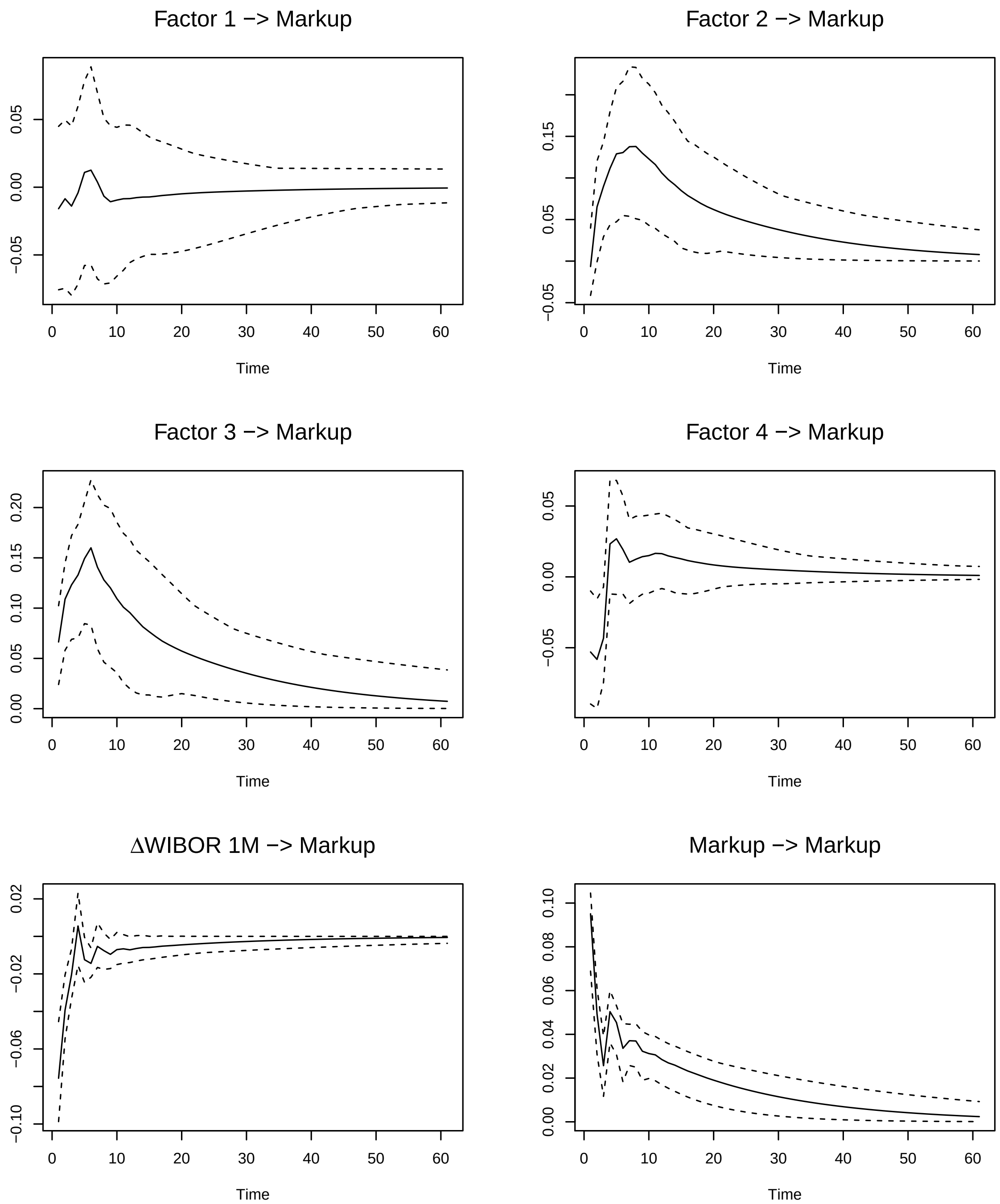


Figure 6: Orthogonalized Impulse Responses of Δ WIBOR 1M

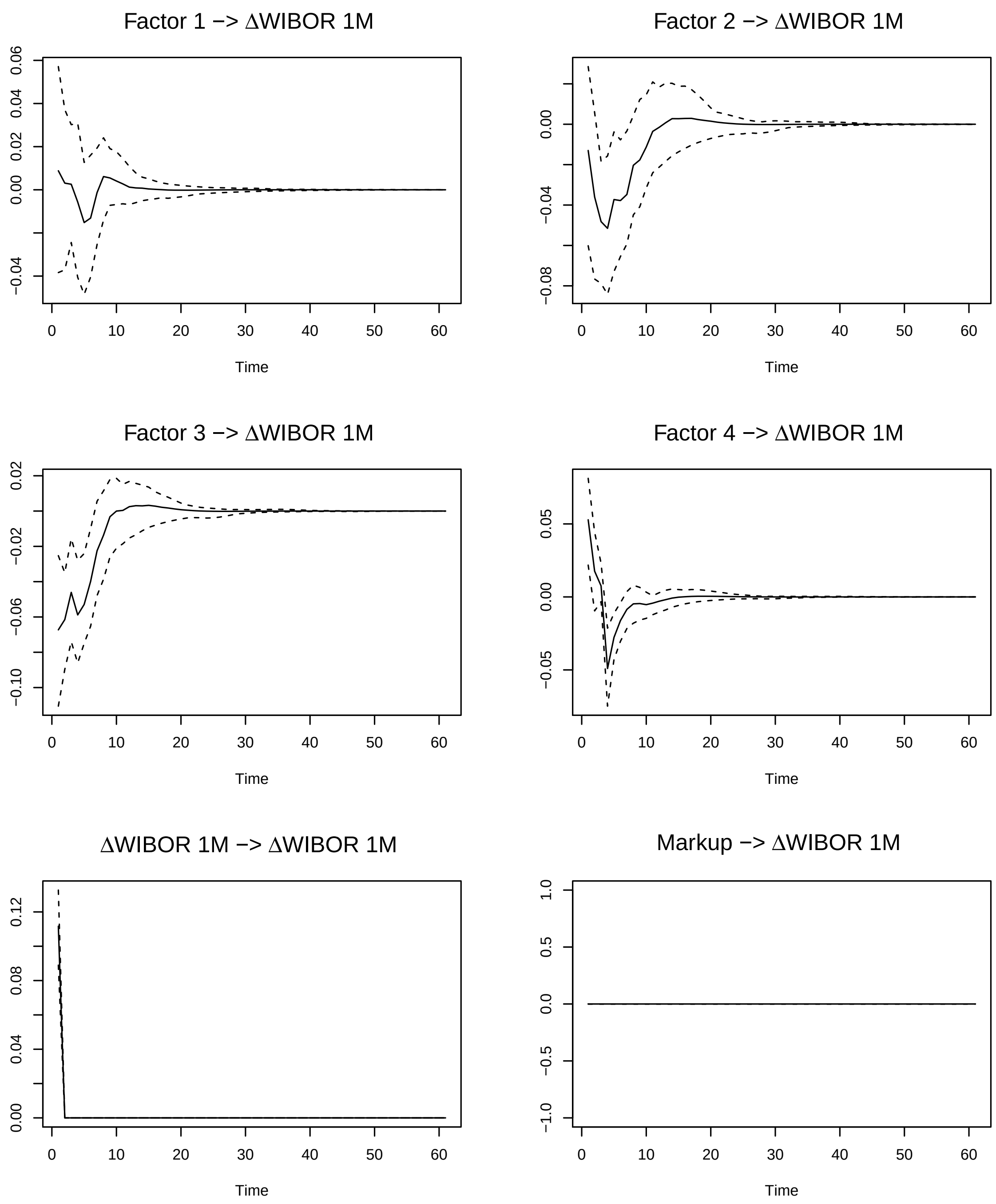


Table 4: Stability Tests, 5-percent level of significance

Andrews Test

Equation	Test Statistic	Bootstrap Critical Value	Asymptotic Critical Value
F_{1t}	8.348	14.126	20.630
F_{2t}	5.973	8.444	15.340
F_{3t}	2.918	8.581	15.340
F_{4t}	5.544	10.112	17.250
Δm_t	8.075	12.193	19.070
z_t	16.380	21.520	24.310
All	36.318	43.461	NA

Andrews-Ploberger Test

Equation	Test Statistic	Bootstrap Critical Value	Asymptotic Critical Value
F_{1t}	1.813	4.918	7.490
F_{2t}	2.096	2.504	5.110
F_{3t}	0.732	2.622	5.110
F_{4t}	1.595	3.350	5.960
Δm_t	2.442	4.117	6.790
z_t	5.612	7.935	9.200
All	14.985	18.750	NA