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CONTENTS

Foreword

On April 23-24, 2015, the National Bank of the Republic of Macedonia organized the 4th Annual Research Conference entitled: "Addressing structural rigidities in view of monetary policy transmission effectiveness". This Conference is traditionally organized on the occasion of the anniversary of the monetary independence of the Republic of Macedonia. The 4th Annual Conference started with a panel of the Governors of the Central Banks of several countries of the region on the main Conference topic. In addition, high quality papers were presented, received upon Call for papers sent to the central banks in the region. This booklet incorporates the papers presented at the Conference, as well as the official speech of the Governor of the National Bank of the Republic of Macedonia.

By launching this booklet, we would like to express our gratitude to the esteemed Governors, all presenters, the discussants, the moderators of the Conference sessions, as well as to all other participants, who all together added value to the success of the Conference.

National Bank of the Republic of Macedonia

Dear Governors, Your Excellences,

Ladies and Gentlemen, Dear guests,

It is my pleasure to welcome you all, on our fourth Research Conference, entitled Addressing Structural Rigidities in View of Monetary Policy Transmission Effectiveness. Just like the previous years, we hope that the topic of this conference is both interesting and challenging to provide an interactive framework for discussion and research-based policy advices for the central bankers.

Structural rigidities, indeed, are related to a wide range of economic segments. Therefore, today we will try to tackle different economic issues, enlightening their importance for the monetary policy transmission effectiveness. Transition economies have undergone many structural changes in the ongoing reshaping of the economies, which to some extent have influenced the monetary policy decision making or monetary policy transmission channels. Globalization is yet another factor that tends to produce changes in the economic structure and at the same time, affects the markets functioning or economic agents' behavior. The labor market rigidities, wage and price settings, financial market imperfection, impact of foreign direct investments and many other issues must be appropriately taken into consideration in the monetary policy implementation and probably we have all put in a great deal of effort to accommodate our models to capture these specifics.

The last global crisis, with its overall severity, time length and countries coverage reactivated the importance of structural issues for the advanced economies too, pointing to the fact that not always same measure fits any situation, not always same risk affects in a same way and not always same threshold is valid for everyone. Nowadays, we speak about flexible and innovative monetary policy, designed on a broadly based information platform, incorporating structural features of the economies and all specifics of the monetary policy transmission mechanism. Since the onset of the crisis, we have tried to better incorporate financial stability aspects into monetary policy decision making. In addition, we learnt that zero lower bound rate in the advanced economies or historically lowest interest rates in emerging economies were not able to revive credit growth, until we moved towards renovating our toolkit. Non-standard monetary policy measures in different modalities have been widely used by the central banks around the globe during the crisis period, with the ultimate goal to capture countries' specific structural features. These tailor made measures proved to have ability for more effective monetary signals transmission, considering the additional impairments of the traditional transmission channels during the crisis.

Taking care about financial stability risks when conducting monetary policy, is one of the most important lessons learnt from the global crisis. In the highly globalized world, in practice, it means that we need to follow interlinks between domestic banking system and international financial markets in view of financing structure as well as its risk exposure considering traditional versus complex market products. Countries of the region, with relatively large share of foreign ownership in the banking system, probably to some extent experienced certain deleveraging pressure on the group level in the past years that unavoidably affected domestic subsidiaries' behavior. Strengthening of regulatory requirements and their implementation in practice, as well as restructuring processes of the EU banks hit by the crisis, impose additional challenges to the financial stability and the monetary policy, as well.

On the backdrop of the lessons from the crisis and still fragile global environment, monetary policy should be designed carefully and implemented flexibly. Besides the traditional transmission channels, we must be aware of the expectations channel as well as the risk-taking channel. Therefore, in implementing our policy measures within the targeting horizon we need to take into account the market expectation as well as the risk perceptions of the economic agents. A mix of standard and non-standard instruments, intensified communication with the public, a broad overview of financial stability risks and the effects of macroprudential regulation, regular monitoring of the external factors' potential spillovers to the domestic economy, are probably the best options for securing effective implementation of the monetary policy.

Although being far away from the acute stage of the crisis, the current global environment is marked with relatively weak and uneven growth outlook, especially in the advanced economies, which led to different paths of their monetary policy stance. While ECB has just started its large-scale quantitative easing program, FED is just about to exit the accommodative monetary policy stance. These contrasting growth prospects of the leading economies underpin the risk of global financial markets volatility in view of eventual rebalancing positions between the main world currencies. On the other hand, there is a group of emerging economies with better growth performances and prospects for the future, including the Macedonian economy, which in the recent years has shown sound recovery, taking into account the significant ongoing structural changes in the economy as well as the fiscal space from the previous years. A common feature for the current surroundings are disinflationary or deflationary pressures around the globe in light of the fall of oil prices, with potential impact on the monetary policy cycle, additionally increasing uncertainty about external environment. At this point, monetary policy must broaden its scrutiny to involve every important aspect from the external environment, domestic economic cycle as well as structural features important for the future monetary policy stance.

The bottom line is that in a dynamic external environment we must take care about structural bottlenecks, regardless of whether they are related to the transitional stage of our economies or induced by the global crisis. I have tried to tackle some of them, which I believe will be discussed further on the high level Governors panel or additionally elaborated in the next sessions within the research papers. As I mentioned at the beginning, this is a broad topic, which is very important for the monetary policy transmission effectiveness. Therefore, I appreciate your attendance and contribution to our Conference, in sharing experiences and broadening our overall knowledge on these issues.

I wish you interesting conference and fruitful discussion!

Dimitar Bogov, Governor of the National Bank of the Republic of Macedonia
Skopje, 23 April 2015

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TRADE IN GOODS, GLOBALIZED PRODUCTION STRUCTURE AND INFLATIONARY DYNAMICS: CROSS COUNTRY EVIDENCES

Draft

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Abstract

This paper empirically assesses the significance of domestic and external factors on the determination of CPI inflation dynamics in a sample of OECD countries and offers evidences on the open economy New Keynesian Phillips Curve (NKPC). Results suggest that globalization in goods market as well as production and financial markets have become more relevant inflation drivers than domestic output gap. Magnitude of the coefficients of each factor changes with respect to time. Final goods trade and globalized production have opposing impacts suggesting that price setting mechanism in final and intermediate goods markets are different, further complicating the monetary policy actions in stabilizing domestic inflation. Signs of both output gap and financial integration change significantly from post-to-pre 2002. Financial integration has begun to play significant decreasing role, while output gap has been generating a pressure on domestic inflation in recent years.

1. Introduction

As globalization expands in different directions including trade in goods, international production fragmentation and financial markets integration, it raises important challenges to monetary economist and policymakers: how and at what extent globalization influences domestic inflation dynamics. Rogoff (2003) argues that globalization played significant supporting role in lowering inflation in the past, in contrast to Ball (2006) who claims no significant relationship between globalization and inflation. The literature provides evidences on both negative (Romer 1993; Lane 1997; Gruben and McLeod 2004; Bowdler and Nunziata 2006); Borio and Filardo 2007), Bowdler 2009) and positive (Deniels et al. 2005; Narayan, et al. 2011) impact of trade openness on inflation, while Razin and Loungani (2007), Badinger (2009), and Bianchi and Civelli (2014) show that both trade and financial openness are significant factors in decreasing inflation. For others, such as Temple (2002), Ball (2003), Sachside et al. (2003), Alfaro (2005), Wu and Lin (2007), and Ghosh (2014) there is no systematic relationship between trade openness and inflation.

In this paper we explore the hypothesis that globalization may decline the effectiveness of monetary policies by empirically investigating potential effects of both trade openness and financial integration. We differ from the existing literature by examining independent impact of trade in final goods and globalized production integration on domestic inflation. Trade in final goods largely includes exchange of substitute goods thus changes the consumption bundle of households and have impulse on domestic consumer prices depending on the strength of pass-through from import prices to domestic prices. Vertically globalized production structure by linking domestic production to international production chains involves intensive trade in intermediate or complementary goods and allows domestic macroeconomic dynamics to become more sensitive to international disturbances of productivity, demand and inflation.

Trade in different forms of goods (consumer goods and intermediate goods due to production chains) may have different impacts on domestic inflation due to possibilities of different price setting strategies of the firms in two different markets (Krugman, 1987; Devereux, 1997). In case of producer-currency pricing (PCP) producers set price of their products at the same currency in all markets, where as in case of local-currency pricing (LCP) firms set prices at the exporting countries currency (See Engel (2009) for detailed analysis). The selection of the pricing strategy depends on the market structure. In this paper we are interested in examining if the nature of trade plays different role in influencing domestic prices complicating the conduct of monetary policy in stabilizing domestic inflation.

We are also interested in at what extend financial integration affects domestic inflation dynamics together with the trade openness. Financial integration expected to influence domestic inflation by easing the flow funds across the countries. According to Tytell and Wei (2004) higher financial integration by increasing elasticity of demand between domestic and foreign currencies could discipline Central Banks to use monetary policy, hence reduce inflation. Existing literature generally provides supporting evidences for financial integration in falling inflation rate (see for instance Gruben and McLeod 2002, Badinger 2009, Ghosh 2014).

We apply panel fixed effect techniques to assess the significance of domestic and external determinants of CPI inflation dynamics in a sample of OECD countries. Panel fixed effect techniques are also used in Ghosh (2014) but this analysis is based on a variant of open economy New Keynesian Phillips Curve (NKPC) derived in Gali and Monacelli (2005) and developed by Wong and Eng (2010) where an indicator for the trade in final goods and for degree of vertical globalized production structure appears explicitly.

In our empirical analysis both domestic factors, trade openness, globalized production structure and financial integration indicators explicitly appear as determinants of inflation dynamics. This study differs from Gali and Monacelli (2005) by accounting for globalized production structure, and it diverse from Wong and Eng (2010) by explicitly pointing out the independent impact of trade openness in final goods, global integration in production and integration in financial markets. Our analysis close to Romer (1993), Lane (1997), Gruben and McLeod (2004), Bowdler 2009) and Deniels et al. (2005), Razin and Loungani (2007), Badinger (2009) , Mihailov et al. (2011a,b) and Ghosh (2014) in a sense that it empirically investigates the impact of trade openness on inflation dynamics, however, unlike them we focus on how different are the impacts of final goods trade and production integration across the countries on domestic inflation dynamics. Our analysis also similar to Razin and Loungani (2007), Badinger(2009), Ghosh (2014) and Bianch, and Civelli (2014) by accounting for financial integration in determination of the inflation dynamics jointly with trade openness. We are different from them not only by accounting for different forms of trade, but also by using interest rate differential (as described in NKPC derivation) as a measure of financial integration. On the measurement of financial openness Gruben and Mcleod (2002), Razin and Loungani (2007) and Ghosh (2014) use capital account restrictions; Denials and VanHoose (2007) use capital inflows and outflows; Badinger (2009), Ghosh (2014) and Bianchi and Civelli (2014) use total assets and liabilities over GDP ratio.

Accounting for the impact of trade openness in final and intermediate goods independently we show that ambiguous empirical evidences on the relationship between trade openness and inflation in the literature hinges on the form of trade openness variable used in the analyses. These studies fail to distinguish the independent impact of trade in goods and production fragmentation on the domestic inflation dynamics. Our estimates suggest that globalization in final goods, production and financial markets have become more relevant inflation drivers than domestic output gap. Though sign and the magnitude of the coefficients of each factor changes with respect to time, final goods trade and globalized production have opposing impacts on domestic inflation. Impact of financial integration is time dependent. Financial integration, unlike pre-2000 period begins to play a significant negative role in decreasing inflation in 2000s.

The reminder of the paper is organized as follows. Section 2 outlines the theoretical background and motivates the empirical strategy. Section 3 sets up empirical model and describes the variables and data used. Section 4 presents the results from the estimation of the benchmark model. Section 5 discusses additional regression analysis. The final section concludes the paper.

2. Theoretical Motivation

Our empirical analysis is based on the model described in Gali and Monacelli (2005) and extended by Wong and Eng (2010). Gali and Monacelli (2005) show that in a small open economy CPI inflation π_t, is determined by domestic inflation and changes in terms of trade:

(1) $\pi_t = \pi_{Ht} + \alpha \Delta s_t$

where $\alpha \in [0,1]$ is index of openness in consumption goods trade (domestically produced goods consumption relative to imported goods), π_t is CPI inflation, π_{Ht} is domestic inflation defined by the rate of change in the index of domestic goods prices, $s_t (= p_{Ft} - p_{Ht})$ denotes the effective terms of trade, p_{Ft} and p_{Ht} are log of foreign and home goods price indexes. The equation indicates that the gap between consumer and domestic price inflation is proportional to the percentage change in terms of trade, with the coefficient of proportionality given by the index of openness. Therefore, both domestic and external factors have a role on the determination of CPI inflation and the impact of external factor depends on the degree of openness.

Dynamics of domestic inflation in terms of real marginal costs (mc_t) are described in Gali and Monacelli (2005) by the following equation:

(2) $\pi_{Ht} = \{\pi_{Ht+1}\} + \lambda mc_t,$

Substituting Eq. (2) into Eq. (1)

(3) $\pi_t = \{\pi_{Ht+1}\} + \lambda mc_t + \alpha \Delta s_t,$

From Eq. (1): $E\pi_{Ht+1} = E\pi_{t+1} - \alpha \Delta s_{t+1}$

Then, Eq. (3) can be arranged as

(4) $\pi_t = \{\pi_{t+1}\} + \lambda mc_t + \alpha(\Delta s_t - \beta \Delta s_{t+1})$

Wong and Eng (2011) show that real marginal costs for a firm producing at the nth stage of a globalized production process at time t using intermediate inputs from abroad and from domestic markets can be expressed in the following form:

(5) $mc_{nt} = \sigma p'_{nt} + (1-\sigma) w_{nt} - a_{nt}$

where p'_{nt} is log of the producer price index for intermediate input, w_{nt} is log of the cost of labor, a_{nt} is log of the state of technology for stage n production. Parameter σ is the share of intermediate input use in production. Producer prices can be computed as the weighted average of the cost of domestic and imported intermediate inputs as,

(6) $p'_{nt} = (1-\kappa) p'_{Hnt} + \kappa p'_{Fnt}$

where $\kappa \epsilon [0,1]$ is an indicator of the extent of globalization in production structure , p'_{Hnt} and $p'_{Fnt}$are logarithm of the price of domestic and foreign intermediate inputs.

Let Qt be the relative price of imported over the domestic goods at the nth stage of the production process, then log of Qt, qt:

(7) $q_t = p'_{Fnt} - p'_{Hnt}$

Substituting Eq. (7) into Eq. (6):

(8) $p'_{nt} = p'_{Hnt} + \kappa q_{nt}$

Substituting Eq. (8) into Eq. (5):

(9) $mc_{nt} = \sigma p'_{Hn-1t-1} + (1-\sigma) w_{nt} - a_{nt} + \sigma\kappa q_{t-1}$

Note that

(10) $y_{nt} = \sigma p'_{Hn-1t-1} + (1-\sigma) w_{nt}$

refers to the income definition of output, y_{nt}, then Eq (9) can be written as

(11) $mc_{nt} = y_{nt} - a_{nt} + \sigma\kappa q_{t-1}$

At the final Nth stage Eq. (11) will be

(12) $mc_t = y_t - a_t + \sigma\kappa q_{t-1}$

Substituting Eq.(12) into Eq. (4)

(13) $\pi_t = \{\pi_{t+1}\} + \lambda(y_t - a_{t)} + \lambda\sigma\kappa q_{t-1} + \alpha(\Delta s_t - \beta \Delta s_{t+1})$

According to the Eq. (13) current inflation is determined by expected future inflation, economic conditions and relative prices in final and intermediate input markets. Changes in terms of trade in final goods market influences current inflation depending on the degree of openness to final goods trade α. Similarly, deterioration in terms of trade in intermediate goods is inflationary to the extent depending on the price stickiness λ and the share of foreign intermediate inputs that indicates the degree of vertical globalized production structure $\sigma\kappa$.

In the next step in order to introduce financial integration into Eq.13 explicitly we use uncovered interest rate parity condition:

(14) $r_t - r_{t*} = \{\Delta_{et+1}\}$

Eq.14 states that under the assumption of complete international markets, the difference between log of domestic bond prices r_t and log of the price of the riskless bonds dominated in foreign currency r_{*t} is equal to the expected changes in log of nominal effective exchange rates. The relationship between terms of trade and nominal effective exchange rates is shown as follows in Gali and Monacelli (2005):

(15) $s_t = e_t + p_{t*} - p_{H,t}$

Combining Eq. (15) and (14) yields the following equation:

(16) $s_t - \{s_{t+1}\} = (r_{t*} - E_t\{\pi_{t+1*}\}) - (r_t - E_t\{\pi_{t+1}\}),$

Perfect foresight assumption implies

(17) $\Delta s_{t+1} = \tilde{r}_{t+1*} - \tilde{r}_{t+1}$

where $\tilde{r}_{t+1*} = r_{t*} - E_t\{\pi_{t+1*}\}$ and $\tilde{r}_{t+1} = r_t - E_t\{\pi_{t+1}\}$. So that, changes in terms of trade are function of real interest rate differential, closer the gap between domestic and foreign rates more integrated the international financial markets.

Combing Eq. (17) and (13) yields an equation for inflation as follows:

(18) $\pi_t = \{\pi_{t+1}\} + \lambda(y_t - a_t) + \lambda\sigma\kappa q_{t-1} + \alpha\Delta s_t - \beta\alpha(\tilde{r}_{t+1*} - \tilde{r}_{t+1}).$

According to Eq.18 domestic current inflation is influenced not only from domestic factors such as expected next period CPI inflation and domestic output gap but also external factors such as relative prices in intermediate goods (q_{t-1}) times the degree of vertically globalized production structure ($\sigma\kappa$), relative prices in consumption goods (s_t) times degree of trade openness in consumption goods (α) and degree of openness in financial markets.

Our empirical analysis is based on Eq. 18 since both domestic and external factors enter explicitly in the determination of domestic inflation dynamics.

3. Empirical Analysis

3.1. Empirical strategy

We use the panel fixed effect techniques to estimate underlying relationship between domestic current inflation and domestic and external factors indicated in Eq.18. Formally, the dynamics of the estimation can be specified in the following econometric modeling:

(19) $\pi_{it} = \theta + \varphi\, x_{it} + \phi y_{it} + \mu_i + \rho_t + \boldsymbol{u}_{it}$

where i represents each country and t time, while φ is a parameter vectors for domestic factors, ϕ is a parameter vector for external factors and u_{it} is a vector of idiosyncratic errors. Domestic stationary variables are included in x_{it}**,** while indicators of stationary external factors are included in y_{it}. Individual country heterogeneity is introduced into the estimations by including a fixed effect parameter $_i$ into the empirical model. The model in Eq. 19 also includes a vector of country-specific time dummies ρ_t to capture aggregate time specific shocks that may affect all countries the same way, such as having an inflation targeting regime or an aggregate productivity shock. Love and Zicchino (2006) suggests removing the time specific effect for each time period by subtracting the cross-sectional mean from each panel member's observation. The time demeaning procedure also eliminates the cross sectional dependence on panel data (Levin et al, 2002). We eliminate time fixed effect dummies before we begin our empirical analysis.

As an empirical strategy first we estimate Eq. 19 by using data covering the whole period and sample of all countries. Then, we follow recursive estimation techniques to examine if there are significant changes in the values of the estimated coefficients. We think recursive estimation is important because our sample period covers 1990-2013 and there are significant policy changes, such as foundation of EU, transition of Central and Eastern European countries, financial turmoil in the USA and Europe during 2008-2009 etc. Also, due to several economic reasons such as regional trade and financial agreements, locational nearness etc. estimated coefficients may vary with respect to the regions, that we are interested in the proceeding section.

3.2. Data and descriptive statistics

We estimate Eq.19 for 34 OECD countries by using unbalanced annual data for the period 1990-2013.[1] In each specification the dependent variable is the annual CPI inflation π_t. In order to compute output gap we follow a standard procedure and take the log deviation of real GDP from a Hodrick-Prescott (HP) trend. The *gap* has additionally normalized by its standard deviation to have comparable magnitudes across the countries.

According to Eq. 19, external factors or globalization influence domestic inflation through increase in trade in goods, production fragmentation and financial integration. Basically there are two approaches to account globalization into the analysis. One may include relative prices (de jure measure) to account for trade policy effects on inflation or actual openness (de facto measure) that is trade relative to GDP to capture impact of trade policy outcomes.[2] Eq. 19 includes both relative prices and actual openness parameters. Therefore, we consider both measures to check the sensitivity of our results to the use of alternative measures. But, we set up our benchmark estimation in line with Temple (2002) argument that predictions of the open economy NKPC are based on importance of trade relative to GDP.

For our analysis it is important to make a distinction between trade in final and intermediate goods. Trade intensity in intermediate goods is used as a proxy for global vertical production structure. Hummels et al. (2001) provide compelling evidence on increased trade in intermediates and intensified global production networks across countries. Trade intensity in final goods measures the degree of flow of substitute final goods across the countries. Our benchmark measure of trade openness in final goods is exports plus imports of consumption goods as share of GDP, TI^{conm}. Similarly trade openness in intermediate goods is the share of sum of exports and imports of intermediate goods over GDP, TI^{intem}. We also check how our estimated coefficients vary if we were using total trade TI^{total}, rather that making distinction between consumption and intermediate goods trade. In addition we check how our results may vary if we use relative prices in consumption TOT^{conm}, intermediate TOT^{intem} and total trade TOT^{total}, measured as a log difference of exports and imports prices as in Mihailov et al. (2011 a,b).

According to Eq. 19 financial integration can be measured as a deviation of domestic bond prices from the world rate FO^{r-r^*}. Lower the deviation higher is the integration. In our benchmark model we use short-term real interest rates taken from OECD data base.[3],[4] As a proxy for world rate we use average of short term interest rates for Germany, Japan and the USA. We also check our results for using alternative de facto measure for financial integration, namely total foreign assets and liabilities as a share of GDP $FO^{funflow}$, Lane and Milesi-Ferretti (2006) measure for financial integration, as in Badinger (2009), Bianchi and Civelli (20014) and Ghoch (2014).

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[1] Australia, Austria, Belgium, Canada, Chile, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Israel, Italy, Japan, Korea, Luxembourg, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Slovenia, Spain, Sweden, Switzerland, Turkey, United Kingdom, United States.

[2] According to Badinger (2009) relative prices are de jure measures of integration and can be used to account for trade policies on inflation, while use of actual openness which is a de facto measure of integration captures the effect of trade policy outcomes.

[3] Nominal short term interest rates are deflated by CPI inflation.

[4] There are missing short term interest rate values for few countries. These values are filled by corresponding rates found in either IFS or Eurostat. For instance for Turkey all rates are taken from Eurostat, for Greece 1990-2000 period rates are taken from IFS, for Japan rates for the period 1990-2002 are taken from Eurostat, for Mexico 1990-1996 observations are taken from IFS and finally for Slovenia 1999-2001 rates are taken from IFS.

Table 1: Descriptive Statistics

Variable	Obs	Mean	Std. Dev	Min	Max
π	762	0.050	0.102	-0.045	1.052
gap	762	0.000	0.010	-0.029	0.029
TI^{total}	762	0.624	0.339	0.109	1.861
TI^{interm}	762	0.422	0.241	0.072	1.280
TI^{conm}	762	0.102	0.052	0.010	0.381
TOT^{total}	280	-0.017	0.044	-0.228	0.081
TOT^{intem}	280	-0.027	0.066	-0.223	0.120
TOT^{conm}	280	-0.009	0.049	-0.163	0.133
FO^{r_r*}	759	0.532	3.098	-21.779	35.118
$FO^{funflow}$	694	6.782	27.169	0.349	240.749

Graph 1: Plot of the key variables (cross sectional average)

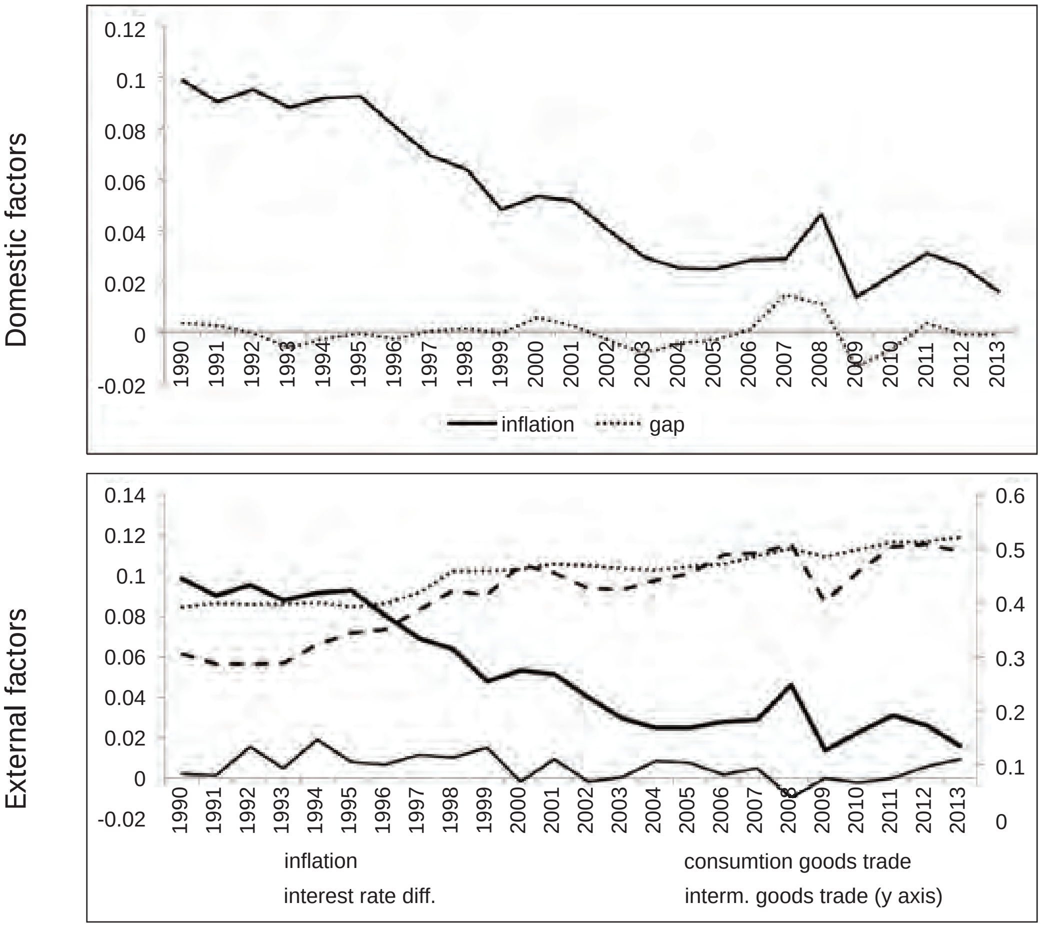


Table 1 summarizes the statistical properties of the variables used in this analysis. There is a considerable variation in the average inflation rate, ranking from -0.05 to 1.05. The same is true for other variables. The highest variations are registered for financial integration indicators. Trade openness indicators measured with trade to GDP ratio exhibit lower variation than relative prices. Graph 1, meanwhile, shows that average inflation rate decreases from 1990 to 20013 while there is an increasing trend in trade related indicators. The average output gap does not show noticeable trend till 2000s but in 2000s it fluctuates in line with the fluctuations in inflation. Just an opposit relationship is observed between inflation and financial integration measured with interest rate differential in 2000s.

4. Estimation Results

4.1. Benchmark analysis

We begin our empirical analysis with the estimation of a version of NKPC which is a pure forward-looking model.[5] Mazinotto (2009) states that an empirical model should include lagged dependent variable inflation to account for the fact that a proportion of firms has backward-looking expectations. Gali and Gertler (1999) and Gali et al. (2001) also, suggest theoretically that firms have to take lagged prices into account just because of the fragmented nature of the production process they are involved. Accordingly, we follow with the analysis of the extended version of NKPC by first including backward-looking inflation expectations, and then adding other explanatory variables one by one to the model. Results are presented in Table 2. Comparison of the results for the model 1 and 2 suggests that both expected inflation and previous year inflation are highly significant with an almost equal coefficient of around 0.5, in determining current year inflation dynamics. Inclusion of lagged depended variable also removes the problem of unexpected negative coefficient for the output gap. But it seems that when the analysis is conducted by including whole period the coefficient for the gap is still insignificant.

Model 3, 4 and 5 compares the closed (model 2) and open economy version of the NKPC. In model 3 we included overall international trade openness; in model 4 we distinguish international openness in consumption goods and intermediate goods; in model 5 we include financial openness. In all estimations the size of the coefficients on expected and lagged inflation as well as output gap is confirmed. Overall, results approve that trade as well as financial openness have significant impacts on current inflation dynamics; therefore they are more relevant factors in determination of inflation than the domestic output gap. Estimated coefficient for financial openness is positive and significant at 5% significance level, suggesting that domestic inflation increases with the increase in interest rate differential or financial disintegration, in line with the findings in the literature (see for instance Badinger, 2009; Razin and Loungani, 2007; Ghosh 2014; Gruben and McLeod 2002).

Table 2: Estimates of NKPC: 1990-2013

Independent variables: inflation (lag and lead), gap, trade openness and financial openness

	Models				
	1	2	3	4	5
(1)	**0.92** 0.02	**0.51** 0.02	**0.51** 0.02	**0.51** 0.02	**0.50** 0.02
(-1)		**0.48** 0.02	**0.49** 0.02	**0.49** 0.02	**0.50** 0.02
gap	-0.12 0.06	0.02 0.04	0.00 0.04	-0.01 0.04	0.00 0.05
TI^{total}			**0.02** 0.01		
TI^{interm}				**0.06** 0.02	**0.06** 0.02
TI^{conm}				-0.17 0.10	-0.16 0.10
FO^{r_r*}					**0.06** 0.03
No. of obs.	728	728	694	694	694
No. of country	34	34	34	34	34

Notes: Bold numbers denotes the coefficients significant at 5% level. Standard errors are reported next to the coefficients

Models 4 and 5 clearly suggest that trade in consumption and intermediate goods have opposing impacts on inflation dynamics. While openness in intermediary goods trade or production fragmentation tends to increase current inflation, trade openness in consumption goods tends to decrease or leave it as it is. This significates the different pricing behavior across the exporting firms in consumption and intermediary goods markets. PCP strategy is expected to be dominant if producers have power to set prices of their own goods which much likely when trade is based on production fragmentation. In these types of international linkages countries are trading intermediate goods for which there is no close substitute in domestic market. LCP strategy may be plausible when there are sufficient substitute goods in domestic market so that exporter firms have to adjust their prices according to domestic market

5 Before beginning with the estimation of Eq. 19 we confirmed that all variables are stationary (TItotal is trend stationary).

conditions, and if necessary they willing to decrease their markups in order not to lose their market. This case is much likely for the consumer goods markets.

Our results suggest that coefficients for trade in consumption goods and trade in intermediate goods have opposing signs, which explains why there is no consensus on the impact of trade on domestic inflation. If trade is dominated by flow of substitute goods it is likely to have negative relationship between inflation and overall trade openness. If, however, trade is dominated by flow of complement goods then it is possible to have positive impact from trade to domestic inflation.

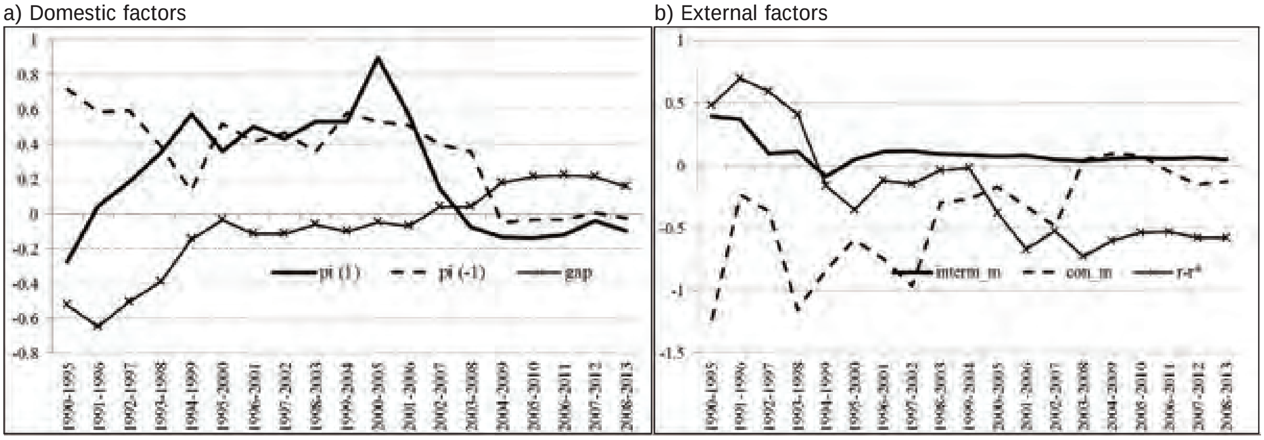
4.2. Recursive Estimates

The period we are analyzing witnessed several economic structural changes, such as foundation of European Union, transition of Central European Countries to the Union, financial turmoil etc. To account for such structural changes here in this section we conduct recursive estimation techniques to examine if there are systematic changes in magnitude and sign of the estimated coefficients in section 4.1.

Graph 2: Path of the Estimated Coefficients from Recursive Regression Analysis

Independent variables: inflation (lag and lead), gap, trade openness and interest rate difference

a) Domestic factors b) External factors



Graph 1 plots the path of the estimated coefficients for domestic and external factors. It is clear that not only the size but also sign of the coefficients for domestic and external factors changes roughly from pre-2000 to post-2000s. In particular, after 2002 coefficient of output gap gains a positive value while those for inflation tend to zero. In case of global factors, none of the coefficients exhibit a stable pattern, except for trade openness in intermediate goods. Financial integration tends to contribute more negatively to the domestic inflation.

Table 3: Estimates of NKPC: 1990-2000; 2001-2013, 2002-2013

Independent variables: inflation (lag and lead), gap, trade openness and interest rate difference

	1990-2000		2001-2013		2002-2013	
	6		7		8	
$\pi(1)$	0.47	0.04	0.55	0.04	0.24	0.05
$\pi(-1)$	0.50	0.05	0.57	0.03	0.53	0.02
gap	-0.32	0.11	0.11	0.04	0.15	0.03
TI^{interm}	0.09	0.04	0.06	0.02	0.07	0.01
TI^{conm}	-0.48	0.30	-0.05	0.09	-0.17	0.08
$FO^{r\_r^*}$	0.22	0.06	-0.24	0.04	-0.53	0.05
No. of obs	284		408		340	
No. of countries	34		34		34	

Notes: Bold numbers denotes the coefficients significant at 5% level. Standard errors are reported next to the coefficients

Results presented in Table 3 note significant changes in the coefficients from 1990s to 2000s. First of all, importance of the domestic factors decreases over time. Current inflation becomes more backward oriented as the coefficient of output gap gains "expected" sign. This finding contradicts with Mihailov et al. (2011a), Borio and Filado (2007) and White (2008) who note a slightly reducing sensitivity of CPI inflation dynamics to domestic output gap.

Recursive path of the estimated coefficient for production fragmentation does not show a significant variation. Comparison of the estimated coefficients for consumption goods trade reveals that substitute goods trade becomes a significant inflation reducing factor in 2000s even though the magnitude of the coefficient falls. Size of the impact of the substitute goods trade found to be larger than that of the intermediate goods trade.

The relationship between financial openness and domestic inflation is not stable and systematic. Financial integration represented with the fall in interest rate difference, supported the falling inflation during 1990s, but in 2000s, the sign of the coefficient is estimated to be negative indicating that financial disintegration in fact is favored for reducing inflation. This result is consistent if we consider the financial turmoil experienced in the USA and European developed countries.

Overall, comparing the absolute values of the estimated coefficients, we may conclude that over the time impact of domestic factors falls while external or global factors become more relevant in inflation dynamics. Contrary to the intermediate goods trade, trade in consumption goods reduces domestic inflation. Impact of financial openness depends on the state of the financial conditions. During turmoil period financial integration could be inflationary.

5. Additional Analysis

5.1. Analysis with Geographical Regions

Are there any differences in how openness influences inflation across country groups? This section focuses on this question and estimate Eq. 19 for different country groups composed of the sample of countries included: Europe, EU15, Central and Eastern European Countries (CEEC) and America.[6] Results presented in Table 4 show that the impact of domestic and external factors changes with respect to the country groups. In general both domestic and external factors take role in inflation dynamics in Europe and in particular in EU15, and the relationship gets stronger in 2000s. On the other hand, for CEEC countries both expected and past inflation are rather relevant for current inflation together with financial integration. For America, overall domestic factors and intermediate goods trade play significant role in inflation dynamics, but in 2000s external factors become more relevant.

6 We do not have sufficient observations for other regional countries to conduct consistent regression analysis.

Table 4: Regional results:

Independent variables: inflation (lag and lead), gap, trade openness and interest rate difference

	Europe		EU15		CEEC		America	
$\pi(1)$	0.50	0.03	0.33	0.04	0.51	0.05	0.45	0.09
$\pi(-1)$	0.49	0.03	0.57	0.03	0.51	0.05	0.59	0.07
gap	0.07	0.05	0.17	0.03	0.06	0.15	-0.29	0.15
TI^{interm}	0.05	0.02	0.03	0.01	0.06	0.05	0.17	0.08
TI^{conm}	-0.12	0.11	-0.09	0.06	-0.21	0.31	-0.52	0.50
$FO^{r\_r^*}$	0.06	0.04	-0.18	0.04	0.16	0.08	0.23	0.12
No. of obs	515		316		126		83	
	1990-2000							
$\pi(1)$	0.48	0.05	0.31	0.08	0.42	0.13	0.43	0.13
$\pi(-1)$	0.48	0.06	0.65	0.06	0.43	0.19	0.65	0.20
gap	-0.22	0.13	-0.01	0.05	-0.55	0.60	-0.62	0.33
TI^{interm}	0.06	0.04	0.02	0.03	0.19	0.15	0.50	0.21
TI^{conm}	-0.40	0.32	0.03	0.10	-3.68	2.22	-5.25	2.13
$FO^{r\_r^*}$	0.22	0.07	-0.20	0.05	0.29	0.22	0.51	0.25
No. of obs	239		151		49		39	
	2002-2013							
$\pi(1)$	0.25	0.05	-0.14	0.04	0.26	0.10	0.15	0.13
$\pi(-1)$	0.55	0.03	-0.07	0.05	0.61	0.04	0.08	0.12
gap	0.15	0.04	0.20	0.03	0.05	0.08	0.06	0.08
TI^{interm}	0.07	0.02	0.06	0.01	0.06	0.03	0.18	0.06
TI^{conm}	-0.18	0.09	-0.17	0.07	-0.17	0.15	0.38	0.34
$FO^{r\_r^*}$	-0.57	0.05	-0.85	0.06	-0.79	0.10	-0.47	0.13
No. of obs	312		180		84		48	
No of countries	26		15		7		4	

Note: Europe: Austria, Belgium, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Israel, Italy, Luxembourg, Netherlands, Norway, Poland, Portugal, Slovak Republic, Slovenia, Spain, Sweden, Switzerland, Turkey, United Kingdom. EU15: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Norway, Portugal, Spain, Sweden. CEEC: Czech Republic, Estonia, Hungary, Poland, Slovak Republic, Slovenia, Turkey. America: Canada, Chile, Mexico, United States. Bold numbers denotes the coefficients significant at 5% level. Standard errors are reported next to the coefficients

These results are consistent with the findings in Mihailov et al. (2011 b) who show that domestic factors tend to play more important role in general in NEUMCs, and Mihailov et al. (2011 a) who support that terms of trade emerges as the more relevant inflation driver than domestic factors largely in small open economies of Europe.[7] However, our results contradict with Ghosh (2014) who finds no clear evidence of a significant effect from trade openness and lower inflation for any country groups.

5.2. Analysis with the use of relative prices

As noted in Badinger (2009), one may use de jure measures of integration to capture impact of economic policy variables such as relative prices, or de facto measures to account for economic policy outcomes. Mihailov et al. (2011a, b) use relative prices in other words log difference of the export prices and the import prices to proxy trade openness with effective terms of trade (TOT). Mazinotto (2009) considers real effective exchange rates as a proxy indicator for international fragmentation, while in Badinger (2009) and Ghosh (2014) measure of openness is the real exports plus imports share in GDP. In earlier section we use de facto trade openness measures for consumption and intermediate goods to conduct our empirical analysis. Here in this section we check how our results may change with respect to the use of alternative trade openness measure, namely TOT computed as a log difference between export and import prices in line with Mihailov et al. (2011a, b).

[7] Countries included in their analysis: Austria, Canada, France, Germany, Italy, The Netherlands, Spain, Sweden, Switzerland and UK.

Table 5: Estimates of NKPC: 2000-2013

Dependent variables: inflation (lead and lag), gap, terms of trade and interest rate difference

	Models					
	1		2		3	
(1)	**0.17**	0.06	**0.17**	0.06	**0.11**	0.05
(-1)	**0.32**	0.05	**0.31**	0.05	**0.33**	0.04
gap	0.31	0.04	0.31	0.04	**0.21**	0.03
TOT^{total}	0.00	0.02				
TOT^{intem}			0.01	0.01	0.02	0.01
TOT^{conm}			0.03	0.02	0.00	0.02
$FO^{r\_r^*}$					**-0.52**	0.06
No. of obs	260		260		260	
No. of country	20		20		20	

Notes: Bold numbers denotes the coefficients significant at 5% level. Standard errors are reported next to the coefficients

We computed TOT for consumption and intermediate goods independently. Export and import prices for these goods are not available for all countries and for the whole period used in the earlier analysis. In present analysis number of countries decrease to 20 and period covered is 2000-2013. Results presented in Tables 5 and 6 suggest that de jure measures of trade openness perform poorer than de facto indicators and are not found as relevant factors for domestic inflation dynamics.

Table 6: Regional results:

Dependent variables: inflation (lead and lag), gap, terms of trade and interest rate difference

	Europe		EU15		CEEC	
(1)	**0.11**	0.05	**-0.11**	0.05	0.21	0.11
(-1)	**0.33**	0.04	-0.03	0.05	**0.49**	0.09
gap	**0.21**	0.03	**0.23**	0.03	0.15	0.09
TOT^{intem}	0.02	0.01	0.01	0.01	0.14	0.08
TOT^{conm}	0.00	0.02	0.02	0.02	0.03	0.05
$FO^{r\_r^*}$	**-0.52**	0.06	**-0.84**	0.05	**-0.48**	0.14
No. of obs	260		195		65	
No. of country	20		15		5	

Notes: Bold numbers denotes the coefficients significant at 5% level. Standard errors are reported next to the coefficients

5.3. Analysis with the use of alternative financial integration indicator

Choice for a measure for financial integration in section 4 is based on the arguments in Eq. 19. Badinger (2009) suggests using an alternative indicator, total foreign assets and liabilities as share of GDP ratio as a proxy variable for financial openness. Ghosh (2014) compares the results from the regression analysis using de jure (capital movement restrictions) and de facto variables. We use the same source referred in Badinger (2009) and Ghosh (2014), namely Lane and Miles-Ferretti (2006) to compute de facto variable to approximate financial integration for our countries. Results from the estimation of Eq.19 by using alternative de facto financial openness measure are presented in Tables 7 and 8. Contrary to the literature we find that inflation is not significantly related to de facto measure of financial openness. It performs poorly compared to the interest rate differentials in the benchmark regression analysis. Comparing the corresponding results for model 5 presented in Table 2 we may conclude that inclusion of de facto measure does not change the parameter values of other explanatory variables. Conclusion does not change if we estimate the Eq. 19 for different country groups (Table 7) or for different time periods (Table 8).

Table 7: Estimates of NKPC

Dependent variables: inflation, gap, terms of trade and de facto measure for financial openness

	All		Europe		Eu15		CEEC		America	
$\pi(1)$	**0.51**	0.02	**0.51**	0.03	**0.35**	0.05	**0.52**	0.05	**0.54**	0.08
$\pi(-1)$	**0.49**	0.02	**0.49**	0.03	**0.53**	0.04	**0.48**	0.05	**0.52**	0.07
gap	0.03	0.09	0.06	0.05	**0.17**	0.04	-0.02	0.15	**-0.55**	-0.32
TI^{interm}	**0.06**	0.02	**0.04**	0.02	**0.04**	0.01	0.06	0.05	**0.19**	0.08
TI^{conm}	-0.20	0.11	-0.17	0.12	-0.08	0.07	-0.31	0.35	-0.68	0.70
$FO^{funflow}$	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.01	0.00	0.01
No. of obs.	660		501		301		121		79	
No. of country	34		26		15		7		4	

Notes: Bold numbers denotes the coefficients significant at 5% level. Standard errors are reported next to the coefficients

Table 8: Estimates of NKPC

Dependent variables: inflation, gap, terms of trade and financial openness

	1990-2000		2001-2013		2002-2013	
$\pi(1)$	**0.46**	0.05	**0.46**	0.04	0.01	0.06
$\pi(-1)$	**0.41**	0.04	**0.55**	0.03	**0.30**	0.03
gap	**0.40**	0.11	**0.15**	0.04	**0.22**	0.04
TI^{interm}	0.07	0.04	**0.07**	0.02	**0.08**	0.02
TI^{conm}	-0.44	0.33	-0.07	0.11	-0.17	0.10
$FO^{funflow}$	0.00	0.00	0.00	0.00	0.00	0.00
No. of obs.	286		374		306	
No. of country	34		34		34	

Notes: Bold numbers denotes the coefficients significant at 5% level. Standard errors are reported next to the coefficients

6. Conclusions

This paper has analyzed the link between globalization and inflation, using data for 34 OECD countries over the period 1990-2013. Globalization we considered comprises both trade and financial openness across the countries. Whist most of the available literature focuses on the overall trade openness as a measure of globalization in goods market, we make distinction between trade in consumption goods and intermediate goods. The incentive for using two different globalization measures is trade in intermediate goods indicates the extent of production fragmentation across the countries so that it measures the intensity of trade in complement goods. Trade in consumption goods largely used to proxy volume of trade in substitute goods. Trade in complement and substitute goods expected to influence inflation in different ways due to the differences in price setting strategies of the firms in these markets.

Our results confirm that external or global factors have become more relevant inflation drivers in recent years. The nature of trade matters in detecting the appropriate trade impact on inflation. While substitute goods trade tends to decrease inflation, trade in complement goods likely to increase it. Financial integration measured with convergence of short term interest rates found to be disinflationary during 1990-2000, but the sign of the coefficient changes significantly for the period 2002-2013, the period covers financial turmoil in Europe and the US.

A further important result is that the predictions of the open economy NKPH largely validated for major European and American countries, the later dominated by the US and Canada. For other countries namely CEEC, which are in their transition period it is hard to detect impact of globalization on domestic inflation.

These results suggest that globalization complicates the conduct of monetary policy. Complexity further increases not only with the heterogeneity in price setting mechanism of firms involving trade in consumption goods and firms integrating to the production networks, but also with the impact of financial integration changing with respect to the economic conditions. As a consequence, we underline the fact that improving competitiveness in both domestic final and intermediate goods markets at different levels from importers to domestic producers could lead to a more effective monetary policy in stabilizing domestic prices. Results one again signifies the stability in financial markets in improving the conduct of monetary policy in controlling domestic inflation.

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PRICE AND WAGE RIGIDITIES IN THE REPUBLIC OF MACEDONIA: SURVEY EVIDENCE FROM MICRO-LEVEL DATA

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This paper exploits the information collected from an ad hoc survey conducted on a sample of Macedonian firms to study the extent of nominal price and wage rigidities in the Republic of Macedonia. The research was motivated by the observation that sticky prices influence the responsiveness of inflation to changes in a central bank's policy rate.

Against this background, the paper investigates the relative importance of most determinants of the frequency of price and wage changes identified in the literature. This paper presents a Bayesian analysis of ordinal data. Posterior inference is carried out using Markov Chain Monte Carlo (MCMC) techniques. Infusing the model with prior information allows us to shrink the parameter space, resulting in more precise and reliable parameter estimates. Our results suggest that higher price flexibility is associated with a higher degree of product market competition. Specifically, we find that firms facing high levels of domestic and international competition tend to adjust prices faster.

JEL Classification: D21, E30, J31
Keywords: price rigidity, wage rigidity, Bayesian inference, MCMC, survey data

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The question of how the price- and wage-setting behavior of firms influences the effects of monetary policy on the economy has taken center stage in the recent literature. Short-run effects of monetary policy on real macroeconomic aggregates are mainly due to the presence of short-term price rigidities, which, through the real interest rate, allow monetary policy to influence real economic activity. Such nominal rigidities play an important role in modern New Keynesian macroeconomic models, which aim to provide key insights on the transmission mechanism of monetary policy to academics and practitioners in central banks and to policy institutions. An understanding of the transmission mechanism is crucial for the correct practical implementation of monetary policy.

Several theoretical studies have outlined the importance of price and wage rigidities on the transmission mechanism of monetary policy (Christoffel et al., 2006) or optimal monetary policy in the presence of wage rigidities (Blanchard and Galí, 2007). Both contributions employ a New Keynesian model with nominal rigidities combined with the Diamond-Mortensen-Pissarides paradigm, thus providing practical guidance on the implementation of optimal monetary policy. This paradigm aims to provide theoretically consistent explanations for phenomena typically occurring in economic systems and their corresponding equilibria. While both studies emphasize the great importance of real rigidity for the actual implementation of monetary policy, the theoretical findings remain inconclusive in explaining how shocks in the labor markets influence monetary policy.

Most theoretical studies provide a rather generic picture, as they investigate aggregate quantities and the reaction of a representative firm to changes of the underlying macroeconomic fundamentals. To provide a deeper understanding of how companies react to shifts in the underlying fundamentals, empirical studies have largely been confined to analyzing individual companies by using large cross-sectional panels. Carlton (1986) and Hall et al. (2000) investigate the pricing behavior of firms facing different degrees of competition. They conclude that firms facing more competition tend to adjust prices faster than companies encountering less competition. Carlton (1986) additionally incorporates the time dimension into the model, extending the analysis by explicitly accounting for persistence effects of demand shocks at some point in time on the price dynamics of a commodity. More precisely, a demand shock today influences not only current prices but also the future path of prices. Geroski (1992) and Álvarez and Hernando (2007) investigate the pricing behavior of firms in different sectors in the U.K. and the euro area, respectively. They corroborate the findings of Carlton (1986) and Hall et al. (2000) and establish that firms operating in less competitive sectors tend to exhibit a somewhat slower reaction to shocks.

This paper investigates the relative influence of several important determinants on the frequency of price changes identified in the literature, such as the degree of product market competition, the cost structure or firms' size. Additionally, we employ a model that is able to track idiosyncratic characteristics and that explains why base wages in some companies tend to be more flexible than in others. These characteristics include the institutional setup for wage bargaining, the composition and characteristics of the workforce, and the wage structure. Using a micro-level survey allows us to unveil the relevance of firm characteristics in the determination of price and wage rigidities, thus enabling us to exploit information that usually cannot be observed in administrative sources. Based on the survey data collected, this paper sheds light on what makes it more or less likely that prices and wages will be sticky, i.e. will not respond immediately to changes in market conditions.

We employ a Bayesian ordered probit model that allows us to incorporate information originating from other studies flexibly and efficiently. Exploiting information from other countries improves the quality of our estimates. Moreover, our Bayesian approach allows us to overcome several problems associated with large numbers of "I don't know" responses and insufficient degrees of freedom. Posterior inference is carried out using the Markov Chain Monte Carlo (MCMC) algorithm put forward by Albert and Chib (1993). In addition, we use a hierarchical prior setup that allows us to set the tightness of the prior in a data-based fashion. This allows us to derive posterior quantities which are infused with prior information when the data become increasingly noninformative.

Our results show that the higher price flexibility is directly related to higher degrees of competitive pressure and exposure to foreign sales as well as to a lower labor cost share. In that respect, our results are consistent e.g. with those of Álvarez and Hernando (2007), who analyze the relationship between price flexibility and competition in nine euro area countries. Our findings are also in line with those of Fabiani et al. (2007) and Vermeulen et al. (2012), who report an inverse relationship between the share of labor cost in total costs and the frequency of price adjustments in nine and six countries of the euro area, respectively. This corroborates the findings in Druant et al. (2009), whose work uses survey data collected in 17 European countries. In addition, the presence of higher workforce turnover, the availability of alternative forms of labor cost adjustment (i.e. of bonuses) along with the presence of any type of wage indexation practice translates into higher wage flexibility. Workforce turnover and the flexible wage component (i.e. the share of bonuses on the firm's total wage bill) are basically margins of adjustment at firms' disposal, in addition to changing base wages, but they could in turn affect wage change mechanisms. Our results are also in line with those of Lebow et al. (2003), Dwyer (2003) and Oyer (2005), who analyze the role of benefits in reducing nominal wage rigidity on the basis of microdata underlying the U.S. Bureau of Labor Statistics' employment cost index (Lebow), Australian microdata (Dwyer), and U.S. data from the National Longitudinal Survey of Youth (Oyer). Their results corroborate those of Druant et al. (2009).

This paper is structured as follows. Section 1 describes the dataset used and provides detailed information on the design of the questionnaire, in parallel presenting some stylized facts emerging from the Macedonian survey evidence in a comparative perspective. Section 2 provides information on the basic econometric framework, prior specifications and the MCMC algorithms employed. Section 3 emphasizes the economic rationale behind the selection of covariates. Section 4 presents the estimation results, and section 5 concludes.

1. Stylized Facts from the Macedonian Survey Evidence Presented in a Comparative Context

The data employed in this paper were collected in a survey which was conducted during the spring of 2014 and which covered a sample of 514 Macedonian firms in manufacturing, construction, trade and other market services. The firms in the final sample account for around 11% of total employment in the Republic of Macedonia. The sample selected is unbiased and representative.[2] The replies seem to be internally consistent. Furthermore, the relatively high response rate (around 80%) promotes confidence in the results. The sample selection is explained in great detail in Ramadani and Naumovski (2014).

The survey applied the harmonized questionnaire of the Wage Dynamics Network (WDN) research project sponsored by a consortium of 23 central banks in the European Union under the lead of the European Central Bank (ECB).[3] This survey was originally carried out by 17 national central banks for countries for which fully harmonized data are available, i.e. Austria, Belgium, the Czech Republic, Estonia, France, Greece, Hungary, Italy, Ireland, Lithuania, Luxembourg, the Netherlands, Poland, Portugal, Slovenia, Slovakia and Spain, between the end of 2007 and mid-2008. The total sample size of the dataset is over 17,000 firms. We use the WDN findings to establish a comparative context for the Macedonian survey evidence discussed below. The WDN has two main research objectives: First, to identify the determinants and features of wage dynamics and labor costs that are pertinent to monetary policy; second, to shed light on the link between wages, labor costs and prices. Furthermore, a series of analytical studies is emerging from this network,[4] thus promoting the circulation of research results and providing a platform for discussion. Among the published research associated with this pooled dataset, we cite Druant et al. (2009), who focus on how European firms' wages and prices are linked, as they provide an infrastructure for our study.

[2] Individual weights were calculated for each firm to make the sample representative of the population of firms and to account for the amount of workers that the firm represents in the population. To this end, three different types of weights were introduced in the dataset: A basic sampling weight to adjust for the unequal probability of firms ending up in the realized sample; an employment-adjusted sampling weight to ensure that the sample represents employees in the population, and a so-called "importance weight" giving each firm in the sample a weight proportional to its size (in terms of employment).

[3] For more details on the WDN survey evidence, please refer to the following link: http://www.ecb.europa.eu/home/html/researcher_wdn.en.html. In addition, the October 2012 issue of Labour Economics 19(5) edited by Etienne Wasmer contains a special section on: Price, Wage and Employment Adjustments in 2007–2008 and Some Inferences for the Current European Crisis.

[4] More information on the pool of research studies arising from this network is available under "Publications" under the following link: http://www.ecb.europa.eu/home/html/researcher_wdn.en.html.

The Macedonian survey questions use 2013 as the reference year. Thus, we find it appropriate to briefly sketch out the prevailing macroeconomic conditions in that period. Economic conditions were broadly favorable in the Republic of Macedonia in 2013. More precisely, following a contraction by 0.4% in 2012, growth accelerated to 2.9% in 2013 and labor markets improved significantly. The recovery was largely driven by the observed broadening of the growth base toward domestic private demand and a better performance of net exports. However, the inflation rate of 2.8% in 2013 to a large extent signaled the transmission of food and import price shocks. In 2013, the financial sector remained resilient. Against this background, monetary conditions were accommodative, with the main policy rate being reduced by 75 basis points to 3.25% in several steps from mid-2012. As a result, credit growth gathered steam from the second half of 2013. However, dynamic household lending growth contrasted with the still weak growth of lending to the corporate sector.

The time gap between the European and the Macedonian surveys spanned the post-2008 global financial and economic crisis period, so that comparisons reflect not only national differences but also changes in the global economic environment. However, note that while favorable economic conditions prevailed in the euro area in the precrisis period, the Republic of Macedonia entered a high-growth period when the survey data were collected.

Several important features of price- and wage-setting behavior have emerged. Below, we focus on some points, in a comparative context, that seem worth emphasizing.

First, the ECB's Final Report of the Wage Dynamics Network[5] (ECB, 2009) shows that prices are adjusted more frequently than wages. This result directly carries over to the Republic of Macedonia: Macedonian survey evidence shows that 30% of the firms revise prices more often than once a year. For the entire euro area, this fraction is 22%, about ten percentage points lower than the non-euro area figure. Moreover, firms that operate in both market services and manufacturing in Macedonia adapt prices much less frequently than those operating in the trade and construction sectors. In parallel, market services have the highest portion of firms reporting that they lack a regular price revision pattern. In addition, in the case of the Republic of Macedonia, survey results show that only 15% of the firms change base wages more often than yearly, which is generally in line with the European aggregate. In this context, around 40% of the European firms confirmed the existence of some correlation between the timing of price and wage changes. Conversely, in the case of the Republic of Macedonia, the majority of firms (70%) did not acknowledge a direct link between the two.

An additional finding stemming from the WDN survey is that wage-setting institutions distinctly determine the nature of both wage dynamics and wage structure. Wage setting displays significant heterogeneity across Europe: Austria, Denmark, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal and Sweden have a broadly regulated system of wage bargaining, which rests on a high number of collective agreements. Conversely, the Czech Republic, Estonia, Hungary, Lithuania, Poland and the U.K. have a largely deregulated system of wage bargaining.

The Republic of Macedonia also uses a broadly deregulated wage negotiation mechanism characterized by relatively loose employment protection. In addition, institutional rigidities are not very strong, social assistance is unlikely to push reservation wages, the tax wedge is modest, and the overall business environment appears to be rather supportive of strong job creation (IMF, 2013). The Macedonian authorities made sizeable efforts to improve the local business environment. Improved indicators raised the Republic of Macedonia's rank to 23rd among the 185 countries in the World Bank's "Ease of Doing Business" index for 2013. To achieve this position, the Republic of Macedonia reduced red tape in a significant number of areas, in turn enhancing working conditions in the private sector most clearly and consequently exerting a positive influence on labor.

The scope to which wages are indexed to inflation in Europe has attracted considerable attention on the part of policymakers. The survey results show that on average, one-third of European firms run a policy that adapts base wages to inflation. Around 29% of the Macedonian firms have a wage indexation mechanism that is predominantly informal and backward looking.

5 The analysis summarized in this report is based on employment-weighted answers. The same type of adjustment is conducted on the Macedonian survey data as well.

2. Econometric Framework

This section provides a brief overview of the modeling framework employed in the empirical application. More specifically, the following subsections describe the general ordered probit model, the prior setup employed and the corresponding posterior distributions.

2.1. The Ordered Probit Model

Following Albert and Chib (1993), we define the vector of ordered responses $Y=Y_1,\ldots,Y_N$ where Y_i takes one of J ordered categories. Moreover, $X=X_1,\ldots,X_N$ denotes a $N \times K$ matrix of exogenous variables. Finally, we define a latent variable Y^*, which is related to Y through the definition of a suitable linking function $F(g)$. Regressing y_i^* on X_i yields the following latent variable model

$$y_i^* = X_i\beta + \varepsilon_i, \varepsilon_i \sim N(0,1) \quad (1)$$

where y_i^* denotes the ith column of Y^* and β is a K-dimensional coefficient vector. X_i is the ith column of X. Conditional on y_i^* equation (1) is a simple regression model which can be analyzed using standard methods. To describe the behavior of y_i^* we introduce a J – dimensional vector $\gamma=(\gamma_0,\ldots,y_J)$ such that

$$y_i = j \text{ if } \gamma_{j-1} < y_i^* \leq \gamma_j \quad (2)$$

where $\gamma_{j-1} < \ldots \leq \gamma_j$ is necessary (but not sufficient) to identify the model.

As mentioned above, the latent variable y_i^* is related to y_i through $F(g)$. Let us denote the probability that $y_i=j$ as $P(y_i=j)$. Under the assumption that $F(g)=\phi(g)$ equals the cumulative distribution function of the standard Normal distribution, the probability of observing $y_i=j$ is given by

$$P(y_i=j|\beta,\gamma) = \phi(\gamma_j - X_i\beta) - \phi(\gamma_{j-1} - X_i\beta) \quad (3).$$

The model described by equations (1) and (3) is not identified. Thus we have to assume that $\gamma_0= - \infty$ and $\gamma_J= - \infty$.

Again, conditional upon knowledge of γ (and thus Y^*), equation (1) reduces to a simple regression model which can be analyzed using standard prior specifications.

2.2. Prior Distributions

Bayesian analysis requires the researcher to specify prior distributions for each coefficient of the model described above. Under the (necessary) assumption that ε_i is standard Normal distributed, we have to choose suitable priors for the elements of β and γ. To control the tightness of the prior on β we introduce a latent hyperparameter $\delta \in R$.

More formally, we impose Normal priors on both coefficient vectors, given by

$$\delta \sim G(a_1,a_2)$$

$$\beta|\delta \sim N(\beta,\delta V_\beta))$$

$$\gamma \propto c$$

The hyperparameter δ is treated as a random quantity, thus it is necessary to impose a prior on δ. We specify a Gamma prior with parameters a_1 and a_2. This choice has several convenient properties because it imposes the restriction that $\delta \in R^+$.

The prior on β is a Normal prior, where $\underline{\beta}$ denotes a K x 1 vector of prior means, $\mathrm{V}_{\underline{\beta}}$ denotes a K x Kprior variance-covariance. This prior, given that the variance of ε_i equals one is conjugate, which facilitates well-known conditional posterior solutions (see Koop, 2012).

Finally, the prior on γ is non-informative and improper for each γ_j. This choice reflects the belief that we have no information on the threshold levels of the latent variable γ_i^* Imposing a diffuse prior on γ, motivated in Albert and Chib (1993) has become a standard choice in the literature on the Bayesian estimation of ordered probit models. Another option would be to impose a Normal prior that fulfills $\gamma_{j-1} \leq \ldots \leq \gamma_J$. However, unless we have strong information on the specific elements of γ, a flat prior proves to be a convenient choice.

2.3. Posterior Distributions and the Markov Chain Monte Carlo Algorithm

Combining likelihood and prior information yields posterior quantities. Under the prior assumptions described above, the conditional posteriors for β,γ and δ take the following form

$$\delta|\beta,\gamma,Y^*,Y \sim p(\delta|Y)$$

$$\beta|\gamma,\delta,Y^*,Y \sim N\,(\overline{\beta},\overline{V}_\beta)$$

$$\gamma|\beta,\gamma,Y^*,Y \sim U(\overline{Y}_{j-1},\overline{Y}_{j+1}).$$

Unfortunately, the conditional posterior of δ is of no well-known form. This fact prevents a simple Gibbs sampling scheme for that parameter. Fortunately, however, the marginal likelihood of the (latent) model in (1) is available in closed form under the conjugate prior. This makes it easy to set up a simple Metropolis Hastings step to simulate δ.

The conditional posterior of β takes a simple form. More specifically, the posterior mean and variance of β is given by

$$\overline{V}_\beta=((\delta V_\beta)^{-1} + X_i^{'}X_i)^{-1}$$

$$\overline{\beta} = \overline{V}_\beta((\delta V_\beta)^{-1}\underline{\beta} + X^{'}y)$$

The latent variable y_i^* can be sampled from the following conditional posterior (see Koop, 2003)

$$y_i^* \mid y_i=j\,,\beta,\gamma \sim N\,(X_i\,\beta,1)\,I\,(\gamma_{J-1} < y_i^* \leq \gamma_J)$$

Where $I(\cdot)$ denotes the heavy side function, which equals one if its argument is true. Thus the posterior of y_i^* is a truncated Normal density from which it is straightforward to sample in general.

Finally, sampling γ can be done quite easily by noting that γ_i has to be between γ_{i-1} and γ_{i+1} Furthermore, we condition on Y and Y^* which implies that we know what value of Y^* corresponds to a given value of Y. This leads to a conditional posterior quantity which is uniformly distributed between $\overline{Y}_{j-1}$ and $\overline{Y}_{j+1}$ (Albert and Chib, 1993).

The conditional posterior distributions described above imply that we can set up a simple Metropolis-within-Gibbs algorithm to simulate the joint posterior of the parameters. Specifically, this implies sequentially drawing the parameters from their conditional distributions with exception of δ, which is sampled through a simple Metropolis step.

3. Data Overview and Prior Implementation

The following section aims at providing a rough overview of the dataset employed and the specifics of the actual prior implementation.

3.1. Data Structure and the Economic Rationale behind the Selection of Covariates

The questionnaire allows us to extend our knowledge of the effects of different labor market institutions and policies on price- and wage-setting schemes. In addition to information on price and wage setting and adjustments, the survey collects data on firms' features, such as the sector of activity, size, structure of the product market, intensity of competitive pressures in the respective market, structure of the labor force and institutional characteristics potentially affecting wage and labor policies.

The dependent variables employed in this paper were constructed as follows. To model price rigidity, a categorical variable was created by breaking down firms' answers to the question on the frequency of price changes. More precisely, the firms were explicitly asked how often they changed the price of their main product. They were able to select one of the following answers: "daily," "weekly," "monthly," "quarterly," "twice a year," "once a year," "every two years," "less than once every two years," "never," or "no predefined pattern." To reduce the complexity, we regrouped the answers into four categories (1 – "daily to monthly," 2 – "quarterly to half-yearly," 3 – "yearly," and 4 – "less frequently than yearly"). Firms that opted for "never" or "no pattern" were not considered in the regression. To model wage rigidity, the value categories of the dependent variable were linked to the degree of stickiness according to one of three categories, with 1 = the firm changes wages more frequently than yearly; 2 = changes wages yearly, and 3 = changes wages less frequently than yearly.

The specific choice of the covariates follows insights provided in Druant et al. (2009), Martins (2013) and Garibaldi (2006). The following section aims to provide a short overview of the explanatory variables included and their economic rationale. The annex provides additional technical information on how the variables were constructed.

The *market competition variable* deduces the degree of competition a firm faces from the relevance it gives to changes in competitors' prices to explain its own price decreases. A firm operating in a more competitive environment and facing higher uncertainty about its future position in the market can be expected to be more concerned with ensuring short-run returns, which leads to higher responsiveness to current shocks.

The external competitive pressure variable is designed to indicate whether prices are stickier when higher portions of a firm's sales are from overseas operations. There is always a tradeoff between the loss of keeping prices unchanged and the cost of adjusting supply. The latter may include fixed costs of entry into the foreign market, which the firm could not recuperate if it decided to scale down supply.

Recent micro-level survey data evidence (see, for instance, Dhyne et al., 2007, Fabiani et al., 2007, and Vermeulen et al., 2012, among others) shows that *labor-intensive sectors* are typically characterized by lower frequencies of price changes, suggesting that stickiness in wages and labor costs may be one of the driving factors behind the slow adjustment of prices.

According to Fabiani et al. (2007), *price reviewing rules* might differ in the presence of frequent shocks: Time-dependent pricing might lead to stickier prices than state-dependent pricing, provided that the time frame is quite large and that the cost of adjustment is low. In the presence of nominal price rigidity, monetary policy can affect economic activity in the short run because it is able to respond to shocks before wages and prices adjust.

The following part of the analysis discusses the logic behind the variables employed as covariates in the nominal wage rigidity model specification.

In an imperfect labor market, *trade unions* play an important role in wage determination. The adoption of a less centralized (i.e. firm-level) wage setting agreement is expected to invoke higher wage flexibility.

The empirical literature points out that *permanent contracts* have a stronger effect on wage rigidity in countries with stricter labor regulations. According to Garibaldi (2006), it is very difficult to measure the degree of enforcement of these regulations because some countries may have rigid standards that are only softly enforced, whereas other apparently flexible countries enforce standards very strictly.

The field literature also suggests that wages of *high-skilled workers* are likely to display higher downward rigidity than those of low-skilled workers. Some characteristics of the labor force might prove to be very important in corroborating this suggestion. For instance, wage compressions (Garibaldi, 2006) could lead to situations in which firms change their recruiting behavior. More specifically, companies could adjust the quantity of their workforce and replace unskilled with skilled workers. The main reason for this willingness to hire overly qualified workers might be the lack of reservations that overly qualified workers will quit as soon as possible, which in turn could be considered an indicator of poor outside options. According to Mojsoska-Blazevski and Kurtishi (2012), overqualification in the Republic of Macedonia is higher than that in most of the EU Member States.

The availability of alternative margins of labor cost adjustment other than adjustment of base wages is essential to evaluate the overall degree of labor cost flexibility. The *share of flexible components* was included to measure the extent to which firms with a higher share of the *flexible pay components* in total labor costs are also those with a lower degree of wage rigidity.

Following the literature, it can be expected that firms experiencing high *workforce turnover* adjust wages more often. A high turnover of skilled workers and a high percentage of novices may be harmful to a company's productivity.

3.2. Prior Implementation

As the harmonized questionnaire of the WDN was used for the Macedonian survey, thus basing the latter on the same underlying theoretical concept as the EU survey, we can exploit information from countries in the EU survey to improve our coefficient estimates. Using the study by Druant et al. (2009) as a reference study, we construct our prior as follows. For the coefficient associated with variable *i*, we center the prior mean $\underline{\beta}_i$ on the corresponding coefficient estimate obtained by Druant et al. (2009). The resulting posterior distribution is thus a weighted average of our data information and the information originating from a study conducted in another country. The weight attached to this specific information is controlled by the hyper parameter δ, which is estimated simultaneously with the other coefficients.

Table 1 - Prior means

Variable	Mean
Price rigidity equation	
Competitive pressure	-0.300
Share of exports	-0.141
Labour-cost share	0.504
State-dependent pricing	-0.241
Wage rigidity equation	
Competitive pressure	0.012
Share of exports	-0.023
Share of permanent workers	0.030
Workforce turnover	-0.170
Share of high-skilled workers	0.012
Collective agreement at firm level	-0.088
Share of bonuses on total wage bill	-0.160
Wage indexation policy	-0.393

Source: Druant et al. (2009).
Note: The data used for this paper consist of a subset of the dataset collected by the Wage Dynamics Network survey. This subset concentrates on 15 EU countries for which fully harmonized data are available, namely Austria, Belgium, the Czech Republic, Estonia, France, Greece, Hungary, Italy, Ireland, Lithuania, the Netherlands, Poland, Portugal, Slovenia and Spain. In addition, the covariates used in our ordered probit models are a subset of our benchmark case, with the exception of the state-dependent pricing variable. The reference for this variable is Martins (2013), who analyses the survey data of Portugal.

The hierarchical nature of our model implies that we let the data inform us about the appropriateness of the prior choice. Thus the question on whether the study by Druant et al. (2009) is appropriate in our context is handled in an automatic fashion. Additionally, we also estimated our models using uninformative priors on the latent regression model. The results obtained were quite similar to the ones obtained from the baseline model described above.

4. Empirical Results

This section investigates the key determinants influencing the frequency of price and wage changes across Macedonian firms within a multivariate framework.

4.1. Investigating the Determinants of Price Changes

A core part of this overview section basically represents a model of the frequency of price changes that accounts for the interaction of a number of firm-level characteristics, such as the degree of market competition, price reviewing rules, as well as the relative importance of labor costs. The variable frequency of price changes is intended to provide a rough measure of the extent of nominal rigidities.

We estimate an ordered probit model in which the dependent variable is the four-category variable defined in section 2. The model also controls for firms' characteristics, such as the sector of activity (manufacturing, construction, trade or business services) or size (in terms of employees: 20 to 49, 50 to 199, 200 or more).

The results summarized in table 2 confirm the presence of some cross-sectional differences in price rigidity between firms. Comparing firms in manufacturing (the reference category) with their counterparts engaged in construction, trade and market services reveals that the former are less prone to leaving the price unchanged for more than one year. The estimates also show that prices are changed less frequently in large firms (firms with more than 20 employees). Conversely, our survey data confirms that higher price flexibility, observed as an increase in the frequency of price adjustment, is more typical of the small firms that perceive strong or severe market competition. In addition, price setting by small companies is found to be more diverse than price setting by larger companies, which most often use markup over cost as their pricing strategy.

Table 2 - Price rigidity: Posterior means and 95% credible sets

Variable	Mean	Percentile 5%	Percentile 95%
Intercept*	2.106	1.824	2.394
Construction*	-0.994	-1.324	-0.656
Trade*	-1.462	-1.701	-1.224
Market services*	-0.385	-0.628	-0.141
20-49*	0.447	0.201	0.692
50-199*	0.449	0.221	0.680
>200*	0.828	0.444	1.208
Competitive pressure*	-0.251	-0.454	-0.048
Share of exports*	-0.034	-0.037	-0.031
Labour-cost share*	0.461	0.283	0.634
State-dependent pricing	0.126	-0.051	0.296

Source: Authors' calculations.
Note: (*) denote statistical significance at 5 percent

Investigation of the specific market structure shows that firms operating in more competitive environments change their prices more frequently. A similar result is also found for the exposure to foreign markets. Thus, companies that increasingly operate abroad tend to adjust prices faster than their purely domestic counterparts. This corroborates the findings of Hall et al. (2000). The results also indicate that price reviewing rules do not seem to have a statistically significant bearing on the frequency of price changes. The results of the analysis of firms' cost structure confirm that a greater share of labor costs in total costs is associated with lower price flexibility, thus suggesting that stickiness in wages and labor costs might be one of the factors behind the slow adjustment of prices.

While the coefficient estimates described above provide a rough picture of the relative importance of several variables for the frequency of price changes, we are ultimately interested in the probability of price changes. We determine this probability by investigating the marginal effects, which establish a relationship between the covariates and the probability of each company to adjust prices.

Table 3 - Marginal effects - price rigidity

			Percentile	
Variable	Probability	Mean	2.5%	97.5%
Construction*				
	Y=1	0.326	0.173	0.480
	Y=2	0.030	-0.033	0.075
	Y=3	-0.158	-0.232	-0.083
	Y=4	-0.198	-0.255	-0.136
Trade*				
	Y=1	0.431	0.334	0.528
	Y=2	0.092	0.041	0.144
	Y=3	-0.184	-0.241	-0.133
	Y=4	-0.339	-0.407	-0.274
Market services*				
	Y=1	0.104	0.024	0.195
	Y=2	0.047	0.014	0.081
	Y=3	-0.048	-0.095	-0.010
	Y=4	-0.104	-0.175	-0.029
20-49*				
	Y=1	-0.092	-0.145	-0.035
	Y=2	-0.079	-0.140	-0.024
	Y=3	0.024	0.005	0.045
	Y=4	0.147	0.047	0.255
50-199*				
	Y=1	-0.097	-0.151	-0.043
	Y=2	-0.077	-0.133	-0.028
	Y=3	0.029	0.010	0.051
	Y=4	0.144	0.055	0.243
>200*				
	Y=1	-0.136	-0.190	-0.076
	Y=2	-0.155	-0.243	-0.063
	Y=3	-0.003	-0.069	0.038
	Y=4	0.294	0.122	0.466
Competitive pressure*				
	Y=1	0.058	0.002	0.111
	Y=2	0.041	0.001	0.087
	Y=3	-0.021	-0.042	-0.001
	Y=4	-0.078	-0.158	-0.002
Share of exports*				
	Y=1	0.904	0.861	0.935
	Y=2	-0.104	-0.144	-0.062
	Y=3	-0.288	-0.345	-0.237
	Y=4	-0.512	-0.580	-0.442
Labour-cost share*				
	Y=1	-0.113	-0.168	-0.061
	Y=2	-0.068	-0.105	-0.035
	Y=3	0.046	0.022	0.075
	Y=4	0.135	0.074	0.199
State-dependent pricing				
	Y=1	-0.031	-0.082	0.021
	Y=2	-0.019	-0.051	0.013
	Y=3	0.013	-0.008	0.037
	Y=4	0.037	-0.025	0.097

Source: Authors' calculations.

The marginal effects summarized in table 3 show that firms operating in the most competitive environments are 7.8% less likely to leave prices unchanged for more than one year and 5.8% more likely to change prices within a one-month period than firms operating in the least competitive environment. The results also indicate that firms with high exposure to foreign markets are 51.2% less likely to leave prices unchanged for more than one year and 90.4% more likely to change prices within a one-month period than firms with the smallest portion of foreign sales.

Controlling for the cost structure indicates that firms with the greatest share of labor costs in total costs are 13.5% more likely to leave prices unchanged for more than one year and 11.3% less likely to change prices within a one-month period than firms with the least labor-intensive processes. Also, firms with more than 200 employees are 29.4% more likely to leave prices unchanged for more than one year and 13.6% less likely to change prices within a one-month period. Moreover, trade firms are 33.9% less likely to leave prices unchanged for more than one year and 43.1% more likely to change prices within one month than manufacturing firms.

4.2. Investigating the Determinants of Wage Changes

In contrast with the evidence found for price rigidity, the results on wage rigidity summarized in table 4 show that the degree of wage flexibility does not differ substantially across sectors, no matter what sector is used as a reference category. This does not hold for the size variable: The degree of wage rigidity seems to decrease in line with firm size. In other words, wage rigidity is more prevalent in small firms than in large firms. We offer the following explanation for the observation that firm size is associated with more price rigidity but less wage rigidity: According to the survey, small firms facing strong or severe competition that are not involved in collective wage agreements tend to absorb input cost shocks mainly by reducing other costs, but also to a large extent by directly adjusting prices. This explains the higher flexibility of small firms' prices. Conversely, big firms tend to absorb input cost shocks predominantly by reducing other costs and by reducing their profit margins, which can be one reason for the higher rigidity of big firms' prices. The fact that big firms have more flexible wages is a signal of higher allocative efficiency, meaning that big firms generally find it easier to absorb shocks or to adjust to structural changes. Furthermore, small firms more often apply a smaller share of flexible wage components, reducing their wage flexibility. Additionally, small firms with low turnover rates (low quit rates) are characterized by stronger wage rigidity. Assuming that firms with low quit rates are those with high turnover costs, such firms have an incentive to avoid wage cuts in order to reduce (costly) job quits. Firm-level collective bargaining does not seem to have a statistically significant impact on wage flexibility.

The results on the flexibility of firms' cost structure and the characteristics of their labor force show that firms in which flexible pay components (i.e. bonuses) account for a greater share of total labor costs exhibit a higher degree of base-wage flexibility. On the other hand, the results demonstrate that the impact of the share of permanent employees on wage flexibility is not statistically significant. The literature also suggests that wages of high-skilled workers are likely to display higher rigidity than those of low-skilled workers. However, table 4 clearly shows that firms with a higher share of high-skilled workers do not display a statistically different attitude toward wage flexibility than firms with low-skilled workers. To some extent, this might reflect the relatively poorer outside options of high-skilled workers as well as their over qualification. On the other hand, the results show that the use of the alternative price margins of labor cost adjustment (like the adoption of bonus schemes) increases wage flexibility.

Table 4 - Wage rigidity: Posterior means and 95% credible sets

Variable	Mean	Percentile 5%	Percentile 95%
Intercept*	1.275	0.926	1.634
Construction	-0.190	-0.531	0.156
Trade	0.068	-0.198	0.334
Market services	0.130	-0.130	0.389
20-49*	-0.462	-0.745	-0.173
50-199*	-0.627	-0.880	-0.376
>200*	-0.537	-0.872	-0.203
Competitive pressure*	0.366	0.163	0.571
Share of exports	0.000	-0.003	0.003
Share of permanent workers	-0.031	-0.232	0.164
Workforce turnover*	-0.006	-0.007	-0.005
Share of high-skilled workers	-0.109	-0.292	0.073
Collective agreement at firm level	0.089	-0.096	0.273
Share of bonuses on total wage bill*	-0.011	-0.015	-0.007
Wage indexation policy*	-0.372	-0.575	-0.169

Source: Authors' calculations.
Note: (*) denote statistical significance at 5 percent

In addition, the marginal effects summarized in table 5 show that firms operating in the most competitive environments are 11.7% more likely to leave wages unchanged for more than one year and 10.3% less likely to change wages more than once a year than firms which operate under the least competitive pressure. Also, firms with the highest workforce turnover are 34% less likely to leave wages unchanged for more than one year and 85.8% more likely to change wages more frequently than yearly than firms with the smallest staff turnover. In addition, firms that adopt indexation strategies are 11.8% less likely to leave wages unchanged for more than one year and 10.6% more likely to change wages more frequently than yearly than firms that do not follow a policy of indexing wages to prices.

Table 5 - Marginal effects - wage rigidity

			Percentile	
Variable	Probability	Mean	2.5%	97.5%
Construction	Y=1	0.057	-0.053	0.190
	Y=2	0.001	-0.034	0.019
	Y=3	-0.058	-0.173	0.079
Trade	Y=1	-0.016	-0.094	0.071
	Y=2	-0.008	-0.045	0.016
	Y=3	0.024	-0.084	0.137
Market services	Y=1	-0.031	-0.105	0.048
	Y=2	-0.014	-0.056	0.011
	Y=3	0.046	-0.059	0.156
20-49*	Y=1	0.139	0.032	0.254
	Y=2	-0.001	-0.044	0.024
	Y=3	-0.138	-0.223	-0.038
50-199*	Y=1	0.187	0.092	0.288
	Y=2	0.000	-0.041	0.031
	Y=3	-0.188	-0.266	-0.105
>200*	Y=1	0.169	0.038	0.313
	Y=2	-0.017	-0.085	0.017
	Y=3	-0.152	-0.241	-0.046
Competitive pressure*	Y=1	-0.103	-0.180	-0.032
	Y=2	-0.014	-0.035	0.005
	Y=3	0.117	0.041	0.190
Share of exports	Y=1	0.005	-0.083	0.106
	Y=2	-0.003	-0.043	0.020
	Y=3	-0.002	-0.118	0.123
Share of permanent workers	Y=1	0.008	-0.054	0.067
	Y=2	0.003	-0.015	0.027
	Y=3	-0.011	-0.093	0.069
Workforce turnover*	Y=1	0.858	0.822	0.889
	Y=2	-0.517	-0.569	-0.463
	Y=3	-0.340	-0.391	-0.291
Share of high-skilled workers	Y=1	0.028	-0.028	0.084
	Y=2	0.009	-0.008	0.032
	Y=3	-0.037	-0.112	0.036
Collective agreement at firm level	Y=1	-0.023	-0.079	0.034
	Y=2	-0.008	-0.031	0.010
	Y=3	0.030	-0.044	0.107
Share of bonuses on total wage bill*	Y=1	0.361	0.187	0.534
	Y=2	-0.103	-0.223	-0.008
	Y=3	-0.258	-0.332	-0.171
Wage indexation policy*	Y=1	0.106	0.034	0.180
	Y=2	0.012	-0.007	0.033
	Y=3	-0.118	-0.189	-0.042

Source: Authors' calculations.
Note: (*) denote statistical significance at 5 percent

5. Conclusions

This paper exploits the information collected from an ad hoc survey conducted on a sample of Macedonian firms to study the extent of nominal price and wage rigidities. The data show that in the Republic of Macedonia, changes in wages occur less frequently than changes in prices. Wages tend to remain unchanged for an average of 16 months. In addition, job tenure is the most important factor behind wage adjustments. Unlike wages, prices tend to remain unchanged for only 7 months. Prices of firms in construction, trade and market services are consistently found to be less sticky than those of firms in manufacturing. The estimates also show that prices tend to be stickier in large firms (firms with 20 or more employees). In addition, unlike price rigidity, the degree of wage flexibility does not differ substantially across sectors. This does not hold for the size variable: Large firms (firms with 20 or more employees) tend to have more flexible wages.

The multivariate analysis of the determinants of price and wage rigidity at the firm level confirms that more frequent price adjustments are associated with more intense competitive pressure and a higher exposure to foreign markets as well as with a lower share of labor costs in total costs.

Higher wage flexibility, on the other hand, is contingent on the presence of higher workforce turnover, the availability of margins of labor cost adjustment other than changes in wages, as well as on the presence of formal or informal wage indexation clauses. The Bayesian approach employed in this paper allows us to combine the prior information obtained from existing studies with our data information, thus effectively updating our beliefs. This mechanism in fact sets the floor for a comparative dimension. Basically, this comparative dimension is built into the model's logic, so that we are able to draw reasonable conclusions about the price and wage rigidity similarities and differences between the Republic of Macedonia and the EU. This framework is rather general and can be employed as a platform for bilateral comparisons between any individual countries or between a country and the average EU outlook.

The survey data are also largely consistent with the macro evidence, notably in the light of macro prudential adjustments to address employment and wage cuts in the aftermath of the global financial and economic crisis. Finally, the inflation outlook in the post crisis period reflects firms' strategies of adjusting prices after facing an adverse demand shock with the intention of counteracting the negative effect of the demand shock as much as possible.

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Annex: Covariates – Technical Summary

Determinants of Price Stickiness

Competitive pressure: dummy that takes a value of one if a firm considers a price decrease likely or very likely when its main competitors decide to cut their prices

Share of exports: export sales as a percentage of total turnover

State-dependent pricing: dummy that takes a value of one for firms that reply that they change their prices without any predefined frequency (prices are reviewed in response to movements in economic conditions) and zero otherwise

Labor cost share: dummy that takes a value of one for firms whose labor cost share overshoots the sample's median share (35%) and zero otherwise

Determinants of Wage Stickiness

Collective agreement at firm level: dummy that takes a value of one if the firm adopts a firm-level collective agreement

Share of permanent workers: dummy that takes a value of one for firms whose share of permanent workers is equal to or greater than the sample median (85%)

Workforce turnover: workers who leave the firm as a percentage of the total workforce (total number of employees in the firm)

Share of high-skilled workers: dummy that takes a value of one for firms in which the share of high-skilled employees is equal to or greater than the sample median (74%)

Share of bonuses on total wage bill: bonus payments as a percentage of total labor costs

Wage indexation policy: dummy that takes a value of one for firms that adopt any form of wage-to-price indexation and zero otherwise

Table A.1: Descriptive statistics

Variable	N	Minimum	Maximum	Mean	Std. Deviation
Competitive pressure	514	0	1	0.72	0.45
Share of exports	514	0	100	24.00	37.56
Labour-cost share	514	0	1	0.49	0.50
State-dependent pricing	514	0	1	0.54	0.50
Share of permanent workers	514	0	1	0.66	0.48
Workforce turnover	514	0	100	24.20	87.09
Share of high-skilled workers	514	0	1	0.55	0.50
Collective agreement at firm level	514	0	1	0.42	0.49
Share of bonuses on total wage bill	514	0	100	10.79	22.47
Wage indexation policy	514	0	1	0.24	0.43
Frequency of price adjustments	329	1	4	2.62	1.05
Frequency of wage adjustments	417	1	3	2.06	0.70

Source: Authors' calculations.
Note: More detailed information on the dataset and the survey used is available on request.

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CENTRAL BANK CREDIBILITY AND THE EXPECTATIONS CHANNEL: EVIDENCE BASED ON A NEW CREDIBILITY INDEX

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Abstract

This article investigates the relationship between central bank credibility and the volatility of the key monetary policy instrument. Two main contributions are proposed. First, we propose a time-varying measure of central bank credibility based on the gap between inflation expectations and the official inflation target. While this new index addresses the main limitations of the existing indicators, it also appears particularly suited to assess the monetary experiences of a large sample of inflation-targeting emerging countries. Second, by means of EGARCH estimations, we formally prove the existence of a negative effect of credibility on the volatility of the short-term interest rate. Thus, in line with the expectations channel of monetary policy, the higher the credibility of the central bank, the lower the need to move its instruments to efficiently fulfill its objective.

Code JEL: E43, E52, E58.

Keywords: Credibility, Inflation targeting, Emerging countries, EGARCH, Expectations.

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1 Introduction

"In a word, credibility matters in the theory and it is certainly believed to matter in practice - although empirical evidence on this point is hard to come by because credibility is not easy to measure" (Blinder (2000, p.1421)). This quotation perfectly sums up the complex issues surrounding the concept of central bank credibility. Credibility is an issue of critical importance in modern central banking (Gonzalez-Paramo (2007)), and is viewed as a precious asset not to be squandered (Blinder (1998)). Nonetheless, despite the growing interest of policymakers and academics in this concept, no clear consensus has emerged on what central bank credibility really means, how it can be established, and especially how it can be measured. The survey conducted by Blinder (2000) indicates that the definition of credibility is not the same for central bankers as it is for academics. In particular, the former more closely relate inflation aversion to credibility than the latter[1].

According to Blinder (1998, 2000), such differences in view between practitioners and academics stems from the fact that the former have a definition of credibility in mind that differs from that formalized within the traditional time-consistency literature originating from Kydland & Prescott (1977) and Barro & Gordon (1983a, 1983b)[2]. Looking back on his experience as a central banker, Blinder (1998) argues that central bankers consider themselves to be credible if their announcements are believed by people, even though they are not bound by a rule that ties their hands and they may have an incentive to renege the initial announcement. In other words, a monetary authority is said to be credible if "people believe it will do what it says" (Blinder (2000)), i.e. if deeds are expected to match words. This short and intuitive definition is close to that considered by Cukierman & Meltzer (1986) in their theoretical work. They define credibility as the absolute value of the difference between the central banks planned monetary policy and the private sectors beliefs about these plans. They define in this way the "average credibility of announcements".

On these grounds, in an inflation-targeting framework, credibility means that people believe that the central bank has the willingness, and also the ability to reach the previously announced inflation target. In particular, this means that private sector inflation expectations are anchored on the target and that people do not over-react to target misses. Based on this statement, several scholars have developed measures for assessing the degree of credibility of a central bank. To the best of our knowledge, the first paper that investigated this issue is Svensson (1993). He compares *ex post* target-consistent real interest rates with market real interest rates on real bonds to assess whether the inflation-targeting framework is credible or not in Canada, New Zealand and Sweden[3]. However, such an indirect approach considers credibility to be a one/zero variable (credible or not, respectively), while in practice, there exist intermediate degrees of credibility (Blinder (1998)). With the increasing availability of survey data on inflation expectations, the next contributions have instead relied on more direct measures of central bank credibility.

Direct credibility measures may be divided into two main categories. The first is based on the Bomfim & Rudebusch (2000) methodology, which consists of assessing the weight attached by the private sector to the inflation target in the formation of their inflation expectations. To this point, if the latter are based on the target, then the central bank is considered to be credible. The second category of central bank credibility measures refers to the gap between the inflation expectations of the private sector and the inflation target (or the inflation target range). The well-known index of Cecchetti & Krause (2002), who define credibility as an inverse function of this gap, belongs to this category. Such an index has been extended by De Mendonca (2007) and De Mendonça & de Guimarâes e Souza (2009), who replace the inflation target point with a target range and consider the possibility of a loss of credibility for negative deviations[4].

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1 Nonetheless, practitioners and academics agree with the reasons why credibility is important, and how to build it. Similar results are obtained by Waller & de Haan (2004) using an updated version of the questionnaire initiated by Blinder (2000).

2 See, notably, Walsh (2010) for an analytical review of this literature.

3 Kupfer (2015) recently used the methodology proposed by Svensson (1993) to assess the monetary policy credibility of the European Central Bank, while Amisano & Tronzano (2010) extended this methodology inside a Bayesian econometric framework.

4 Another category of measures assumes that the current credibility of a central bank is a self-reinforcing process that can be proxied by past inflation performance. In this view, a central bank is expected to gain additional credibility by reaching its publicly announced target repeatedly, i.e., by having "a history of doing what it says it will do" (Blinder (2000)). Considering this assumption, De Mendonca & de Guimaraes e Souza (2009) and Neuenkirch & Tillmann (2014) propose alternative measures of central bank credibility based on the past deviations of inflation from the target. Such backward-looking indicators are particularly relevant for developing countries, for which inflation expectations data are often unavailable.

The indicators developed by Cecchetti & Krause (2002), De Mendonca (2007), and De Mendonca & de Guimaraes e Souza (2009) have two main advantages. They are intuitive and easy to compute. However, they are not discriminating enough. Indeed, they rely on an *ad hoc* parameter for expected inflation, set to 20%, beyond which the credibility of a central bank is considered to be null. Such a threshold is unjustified, considering the single-digit inflation rates and the decreasing inflation targets in the concerned countries over the last decade. When the target is low, these indicators improperly underestimate the effect of large positive deviations of inflation expectations from the target on credibility, particularly when the target is far from 20%.

Against this background, the first purpose of this paper is to propose a new simple time-varying measure of central bank credibility that addresses the main limitation of the existing indexes. Because we believe that in practice, negative deviations of inflation expectations from the target are less likely to compromise credibility than positive deviations, we provide an asymmetric measure of credibility based on the linear exponential (LINEX) function. Furthermore, our indicator does not depend on any *ad hoc* threshold. We compute our index for all emerging countries that adopted an inflation-targeting framework, except for Ghana, for which data on inflation expectations are not available. We then analyze how the credibility of monetary policy has evolved in these economies.

This question of monetary policy credibility is particularly relevant for emerging inflation- targeting countries. A credible central bank is expected to improve the efficiency of monetary policy transmission through two channels: the expectations channel and the interest rate channel. Indeed, if a central bank is credible, people believe that the announced target will be realized. From the observed and expected inflation rate, agents can infer the future path of interest rates. Monetary policy is then easily transmitted along the yield to maturity curve. Moreover, wages and prices are set accordingly. Disinflation is then less costly. Finally, changes in the policy rates are less likely to be considered to be temporary by the banking sector, which is then more prone to pass monetary policy impulses on retail interest rates (Mojon (2001)). At the extreme, the speeches of the governor become an instrument *per se*, and are sufficient for governing the stance of monetary policy. It is not necessary to frequently change the level of key interest rates. Consequently, central bank credibility is a self-reinforcing process that emerging economies should seek to strengthen. For this reason, the second purpose of this paper is to evaluate, in the light on our new indicator, the effect of credibility on interest rate volatility. As far as we know, our study is the first that investigates this issue for a large sample of emerging inflation-targeting countries.

The remainder of the paper is structured as follows. Section 2 provides an overview of the existing measures of central bank credibility. Section 3 presents our new index. Section 4 compares our index with previous indicators and analyzes the evolution of central bank credibility in emerging inflation-targeting countries. Section 5 is devoted to the impact of credibility on interest rate volatility. Section 6 concludes the paper.

2 The existing measures of central bank credibility

Two main types of credibility measures have been developed in the literature. The first refers to the Bomfim & Rudebusch (2000) approach. It consists of assessing the weight the private sector attaches to the inflation target when forming their inflation expectations. More precisely, this approach considers that inflation expectations are determined as a weighted average of the current inflation target and the past inflation rates:

$$\pi^e_{t|T} = \lambda\bar{\pi}_t + (1-\lambda)\tilde{\pi}_{t-q} \qquad (1)$$

with $\pi^e_{t|T}$ representing the inflation expectations of the private sector formed at time t for the period T, $\bar{\pi}_t$ representing the inflation target, and $\tilde{\pi}_{t-q}$ representing the average of past inflation rates over the *q* periods considered $\left(\tilde{\pi}_{t-q} = \frac{\pi_{t-1} + \ldots \pi_{t-q}}{q}\right)$. The parameter λ $(0 < \lambda < 1)$ measures the degree to which expectations are anchored on the target. The higher λ is, the higher the weight attached by the economic agents to the target when forming their expectations, and the higher the central bank credibility. As

Bomfim & Rudebusch (2000) argue, with representative agents, *A* may be interpreted as the subjective probability that an agent attaches to the future achievement of the target. With heterogeneous agents, *A* may represent the fraction of the population believing that the target will be achieved. However, the Bomfim & Rudebusch (2000) approach has received little coverage in the empirical literature, except for the paper of Lysiak, Mackiewicz & Stanislawska (2007) in the case of Poland and those of Demertzis, Marcellino & Viegi (2009) for some industrialized inflation-targeting countries.

The second type of measure refers to the gap between inflation expectations and the inflation target. It considers any deviations of expectations from the target as a loss of central bank credibility. The index developed by Cecchetti & Krause (2002) belongs to this category. Taking values from 0 (no credibility) to 1 (full credibility), it is defined as follows:

$$CRED_{CK} = \begin{cases} 1 & \text{if } \pi^e \leq \bar{\pi}_t \\ 1 - \dfrac{1}{20\% - \bar{\pi}_t}\left[\pi^e - \bar{\pi}_t\right] & \text{if } \bar{\pi}_t < \pi^e < 20\% \\ 0 & \text{if } \pi^e \geqslant 20\% \end{cases} \quad (2)$$

with $\bar{\pi}_t$ representing the inflation target pursued by the central bank and π^e representing the inflation rate expected by the private sector. The central bank is considered to be fully credible ($CRED_{CK} = 1$) if expected annual inflation is lower than or equal to the inflation target. On the contrary, it is non-credible ($CRED_{CK} = 0$) if expected annual inflation is equal to or higher than 20%. Between these two limits, the value of the index decreases linearly as expected inflation increases. This index was first extended by De Mendonca (2007), considering that 1) not only positive, but also negative deviations of inflation expectations from the target can imply a loss of credibility and that 2) in practice the target is not a single value but a range. The following indicator is then suggested by De Mendonça (2007):

$$CRED_{DM} = \begin{cases} 1 & \text{if } \pi^e = \bar{\pi}_t^{mid} \\ 1 - \dfrac{1}{\bar{\pi}_t - \bar{\pi}_t^{mid}}\left[\pi^e - \bar{\pi}_t^{mid}\right] & \text{if } \bar{\pi}_t^{min} < \pi^e < \bar{\pi}_t^{max} \\ 0 & \text{if } \pi^e \geqslant \bar{\pi}_t^{max} \text{ or } \pi^e \leqslant \bar{\pi}_t^{min} \end{cases} \quad (3)$$

with π^e representing the inflation expectations of the private sector $\bar{\pi}_t^{mid}$ representing the midpoint inflation target pursued by the central bank, and $\bar{\pi}_t^{min}$ and $\bar{\pi}_t^{max}$ representing the lower and upper bounds of the inflation target range, respectively. $\bar{\pi}_t$ in the denominator corresponds to the lower bound $\bar{\pi}_t^{min}$ if $\pi^e < \bar{\pi}^{mid}$ and to the upper bound $\bar{\pi}_t^{max}$ if $\pi^e > \bar{\pi}_t^{mid}$ As in Cecchetti & Krause (2002), the index is defined between 0 (no credibility) and 1 (full credibility). While its maximum (full credibility) is obtained when the expected inflation is exactly equal to the midpoint of the inflation range, the index decreases symmetrically and linearly when expectations deviate from the target point.

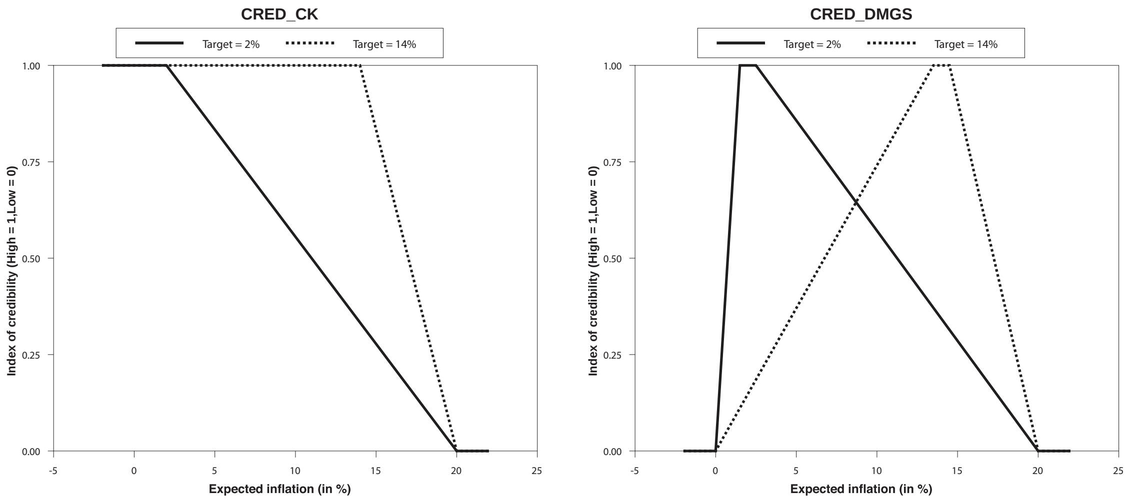
However, in focusing on the midpoint, this index is too restrictive, and can therefore lead to misleading conclusions. Full credibility is not only reached when inflation expectations are exactly equal to the midpoint target. One can reasonably consider that credibility also occurs when private expectations belong to the range. Taking this into consideration, De Mendonça & de Guimarâes e Souza (2009) proposed this alternative index:

$$CRED_{DMGS} = \begin{cases} 1 & \text{if } \bar{\pi}^{min} \leqslant \pi^e \leqslant \bar{\pi}^{max} \\ 1 - \dfrac{1}{20\% - \bar{\pi}_t^{max}}\left[\pi^e - \bar{\pi}_t^{max}\right] & \text{if } \bar{\pi}_t^{max} < \pi^e < 20\% \\ 1 - \dfrac{1}{-\bar{\pi}_t^{min}}\left[\pi^e - \bar{\pi}_t^{min}\right] & \text{if } 0 < \pi^e < \bar{\pi}_t^{min} \\ 0 & \text{if } \pi^e \geqslant 20\% \text{ or } \pi^e \leqslant 0 \end{cases} \quad (4)$$

with π^e representing the inflation rate expected by the private sector and $\bar{\pi}_t^{min}$ and $\bar{\pi}_t^{max}$ representing the lower and the upper bounds of the inflation target range, respectively. A central bank is viewed as non-credible ($CRED_{DMGS}$ = 0) if expected annual inflation is equal or greater than 20% or lower than or equal to 0% and as fully credible ($CRED_{DMGS}$ = 1) if inflation expectations belong to the target range. Between these two limits, the value of the index decreases linearly.

Figure 1 illustrates the profile of $CRED_{CK}$ and $CRED_{DMGS}$ in the case of a single- digit inflation target equal to 2% (with +/- 0.5% point tolerance intervals) and in the case of a double-digit inflation target equal to 14% (with +/- 0.5% point tolerance intervals)[5].

Figure 1: Profile of $CRED_{CK}$ and $CRED_{DMGS}$



As we can see, the profile of these indexes and the marginal loss in credibility largely depends on the level of the inflation target. A positive deviation of inflation expectations from the target is strongly punished in terms of credibility loss if the target is close to the *ad hoc* upper limit of 20%. For example, for a positive deviation of 3 percentage points from the target range, the value of $CRED_{DMGS}$ is then equal to 0.45 in the case of a target equal to 14%, and to 0.83 in the case of a target equal to 2%. The $CRED_{CK}$ index is equal to 0.50 and 0.83, respectively. Such a framework is inadequate for assessing the current level of credibility of emerging inflation-targeting central banks because most of them now pursue relatively low inflation targets. Indeed, in 2014, none of the emerging inflation- targeting countries pursued an inflation target point higher than 5%. Consequently, we propose a new index of central bank credibility, independent of any *ad hoc* upper and/or lower threshold(s). This indicator is described in the next section.

3 A new indicator of central bank credibility

3.1 The rationale for a new indicator

The indicator we suggest is in line with the theoretical considerations of Cukierman & Meltzer (1986), according to which credibility can be viewed as the difference between private inflation expectations and the announced policy target. In this respect, it is an extension of the empirical measures suggested by Cecchetti & Krause (2002), De Mendonca (2007) and De Mendonça & de Guimaraes e Souza (2009). However, as we have seen, these measures impose an *ad hoc* and undue upper threshold value (20%) for expected inflation, above which credibility is null.

We consider that an indicator of credibility should fulfill two main properties. First, it should not be based on *ad hoc* upper and/or lower thresholds but should freely converge towards its extreme values. Second, a credibility indicator should not be linear. Indeed, a critical point for developing a credibility index is the following: should negative and positive deviations of expected inflation from the target

[5] The $CRED_{DM}$ index is not presented here because, as aforementioned, it is certainly too restrictive in that it assumes that credibility is null when inflation expectations are outside the target range.

be considered equivalent in terms of (loss in) credibility? Surely not. The central bank is mandated to maintain control over the growth rate of prices. Positive deviations clearly signal that people do not believe in the ability of the central bank to meet this commitment. Then, the central bank is not entirely credible. Negative deviations also indicate that people believe that actual inflation will not meet the target. However, private agents consider in this case that the monetary authorities can do even better than the announced target in terms of inflation control. This is rarely perceived as a signal that monetary authorities abandon their objective. On the contrary, people consider that "he *who can do more can do less"*[6]. As a result, negative deviations are less serious than positive deviations. An indicator of central bank credibility should take this asymmetry into account, with positive deviations being more serious in terms of credibility loss than negative ones.

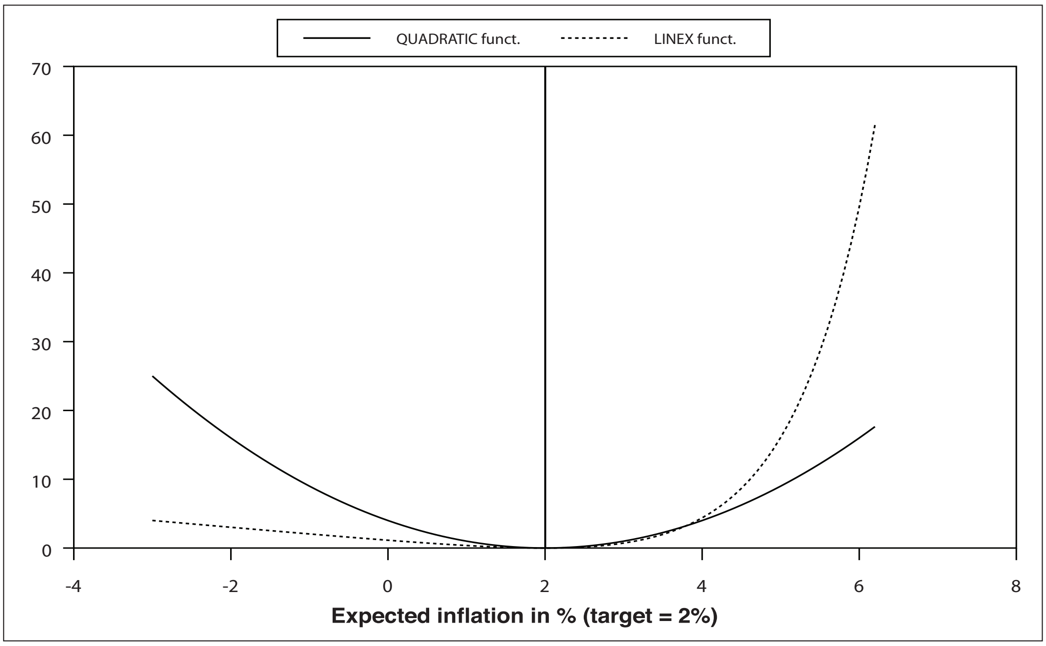
We suggest an indicator that satisfies this dual challenge based on the asymmetrical LINEX loss function[7] (partly LINear, partly EXPonential). Noting *n*[e] the deviation between expected inflation (π^e) and the target $(\bar{\pi})$, a LINEX function with $\tilde{\pi}^e$ as an argument is defined such that:

$$f(\tilde{\pi}^e) = exp(\phi(\tilde{\pi}^e)) - \phi(\tilde{\pi}^e) - 1 \quad (5)$$

For $\phi = 1$, $\tilde{\pi}^e > 0$ will be considered to be more serious than $\tilde{\pi}^e < 0$ (because the exponential part of the function dominates the linear part when the argument is positive). The figure 2 compares the LINEX function with the usual quadratic one for $\bar{\pi}$ = 2%, with the horizontal axis corresponding to π^e.

We will show below that a credibility indicator can be developed on the basis of such a function, with an inverted-U profile between 0 and 1, as is usual in the literature. The indicator will be precisely defined in the next subsections, considering two cases based on whether the target is a single value or a range. Further, in each case, we will successively assume first that negative deviations induce a credibility loss and second that they do not imply any credibility loss.

Figure 2: Quadratic *versus* LINEX functions



6 Typically, the current and expected inflation rates in the euro area have been far below the implicit target of 2%, without impairing the beliefs of agents on the willingness of the European Central Bank (ECB) to control inflation. On the contrary, this achievement reaffirms that European policymakers are perceived to be very attached to their main objective of price stability. Such a situation does not compromise the credibility of the ECB.

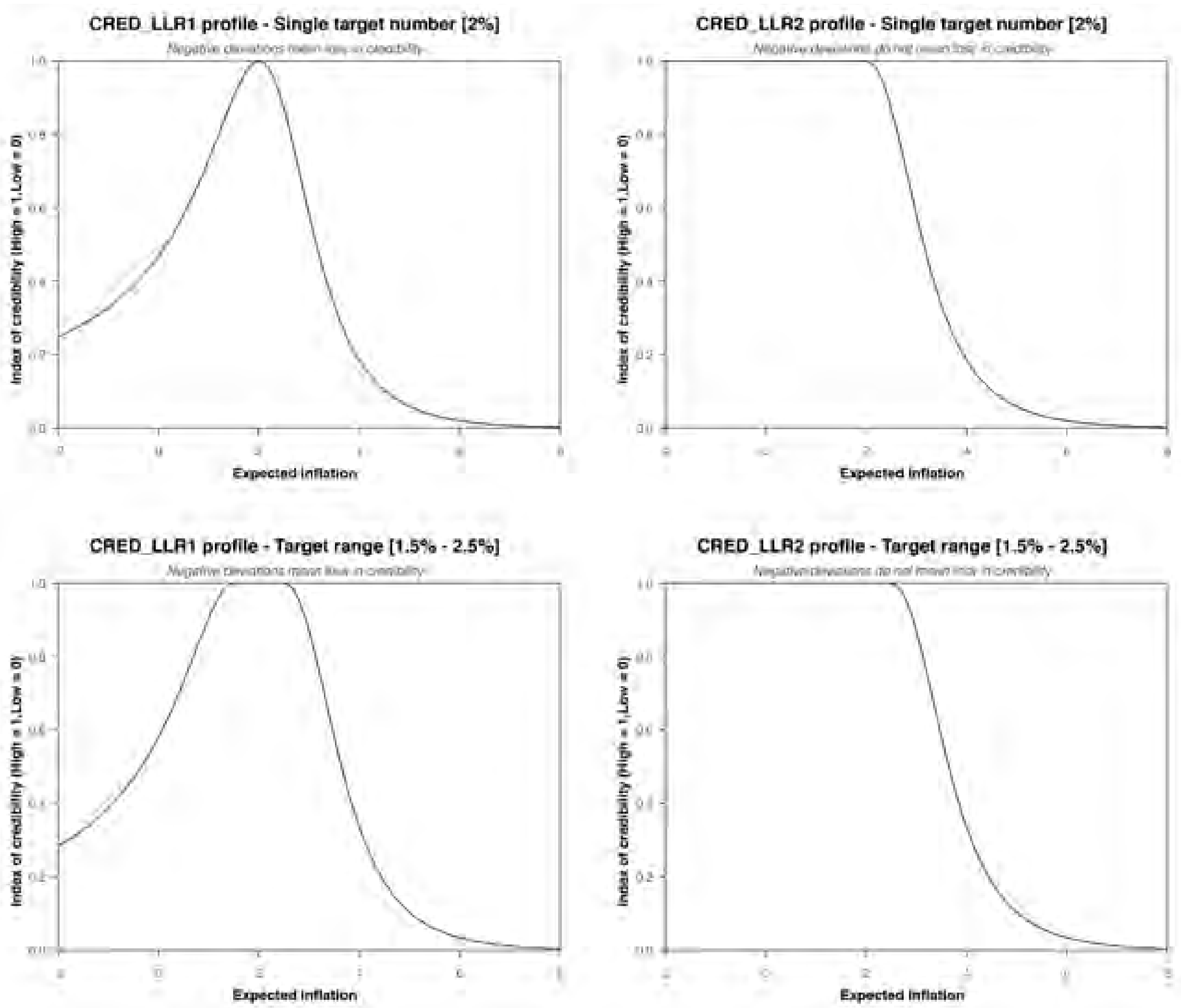
7 See Varian (1974) and Zellner (1986).

3.2 The target is a single number

We first considerer that the target is $\bar{\pi}$.

FIRST CASE One considers π^e < $\bar{\pi}$ to represent a loss in credibility, even if it is less serious than π^e > $\bar{\pi}$. Then, we define a new credibility index as the following inverse *quasi* LINEX function:

$$CRED_{LLR1} = \frac{1}{exp(\tilde{\pi}^e) - \tilde{\pi}^e} \quad \forall \pi^e \tag{6}$$



As for the existing indicators in the literature, 0 < $CRED_{LLR1}$ < 1, with 1 for full credibility. At the extreme opposite, $CRED_{LLR1}$ = 0 indicates that the corresponding central bank is not credible at all. With this definition, any reference to a hypothetical upper or lower bound is not required. The upper left panel of the figure 3 gives the profile of this indicator for $\bar{\pi}$ = 2%, with the horizontal axis representing π^e. As expected, the profile is non-linear. Negative deviations do mean that credibility is compromised, but any positive deviation signals a higher loss in credibility than an equivalent negative one. Moreover, the marginal loss in credibility is decreasing with n^e. This is an important feature of our indicator. The rationale behind is the following. Assume that $\bar{\pi}$ = 2%. An expected inflation rate that grows from 14 to 16% should not coincide with a dramatic loss in credibility, as the latter is already hugely damaged (because of the initial π^e = 14%). Quite the opposite, a growing expected inflation rate, say from 2 to 5%, must express a higher marginal loss in credibility. An inverted-U credibility curve, with a higher slope in the neighborhood of the target than at its extremities, is then justified.

SECOND CASE Considers that $\pi^e < \bar{\pi}$ does not mean loss in credibility. This is an extreme interpretation of the "*he who can do more can do less*" hypothesis. Then, our new indicator simply becomes:

$$CRED_{LLR2} = \begin{cases} 1 & \text{for } \pi^e < \bar{\pi} \\ \dfrac{1}{exp(\bar{\pi}^e) - \bar{\pi}^e} & \text{for } \pi^e \geqslant \bar{\pi} \end{cases} \tag{7}$$

The profile of this credibility function is represented at the top right panel of the figure 3.

3.3 The target is a range, such that $\bar{\pi} = \left[\bar{\pi}^{min}, \bar{\pi}^{max}\right]$

Again, two cases are to be considered, depending now on whether $\pi_t^e < \bar{\pi}^{min}$ is synonymous with a loss in credibility or not.

FIRST CASE $\pi_t^e < \bar{\pi}^{min}$ signals loss in credibility. Then,

$$CRED_{LLR1} = \begin{cases} \dfrac{1}{exp(\pi^e - \bar{\pi}^{min}) - (\pi^e - \bar{\pi}^{min})} & \text{for } \pi^e < \bar{\pi}^{min} \\ 1 & \text{for } \pi^e \in \left[\bar{\pi}^{min}, \bar{\pi}^{max}\right] \\ \dfrac{1}{exp(\pi^e - \bar{\pi}^{max}) - (\pi^e - \bar{\pi}^{max})} & \text{for } \pi^e > \bar{\pi}^{max} \end{cases} \tag{8}$$

The bottom left panel of the figure 3 illustrates this case for a range corresponding to [1.5% — 2.5%].

SECOND CASE $\pi^e < \bar{\pi}^{min}$ does not imply loss in credibility. Then,

$$CRED_{LLR2} = \begin{cases} 1 & \text{for } \pi^e \leqslant \bar{\pi}^{max} \\ \dfrac{1}{exp(\pi^e - \bar{\pi}^{max}) - (\pi^e - \bar{\pi}^{max})} & \text{otherwise} \end{cases} \tag{9}$$

The corresponding profile is represented in the bottom right panel of the figure 3.

4 Application to emerging inflation-targeting countries

$CRED_{LLR1}$ and $CRED_{LLR2}$ are computed for all emerging economies that adopted an inflation-targeting framework except Ghana, for which survey data on inflation expectations are not available. This monetary policy strategy is currently led by 18 emerging countries, while Slovakia abandoned inflation targeting in January 2009 to join the euro area. Our sample is then composed of Brazil, Chile, Colombia, the Czech Republic, Guatemala, Hungary, Indonesia, Israel, Mexico, Peru, the Philippines, Poland, Romania, Slovakia, South Africa, South Korea, Thailand and Turkey.

4.1 Data and periods

For each country, $CRED_{LLR1}$ and $CRED_{LLR2}$ indexes are computed on a monthly basis and cover the period between the effective inflation targeting adoption date (if data on inflation expectations are available) and December 2013. Table 1 provides some details concerning inflation targeting adoption dates and data availability.

Concerning private sector inflation expectations, we use the forecast survey dataset provided by Consensus Economics, which gathers professional analysts forecasts for a large range of macroeconomic variables. The surveyed forecasters are located in their respective country and are working in the financial sector. Therefore, they have a good idea of how inflation will evolve in the medium-term. Moreover, they are more forward-looking than other categories of the population, such as consumers[8]. Because the forecasts are provided for the current and the next calendar year on a monthly basis, we construct a monthly sample of twelve-month ahead expected inflation by taking the weighted arithmetic average of the mean forecast for the current year and the next year, defined as follows:

$$\pi^{e}_{t,12m} = \frac{(12-t)\pi_t^{e^{current}} + t\,\pi_t^{e^{next}}}{12} \tag{10}$$

with t representing the month (with 1 (= January) < t <12 (=December)) at the time of the forecast. Thus, by December, the forecast for the current year is already irrelevant and the forecast for the next year receives full weight *(t*=12). Most of the studies using data from the Consensus Economics adopt this approach for constructing twelve-month ahead forecasts (see, e.g., Beck (2001)). Some emerging countries were surveyed only once every two months at the start of the Consensus Economics survey. For Central and Eastern European countries, surveys have been conducted each month only since May 2007 (see table 1). Linear interpolation can be applied, but cautiously, given the particularity of the data. Indeed, if the missing observation refers to December or January, the interpolated data will overlap two different years. This is a problem when data aims to measure the expected evolution of the consumer price index over a given year. So, we distinguish two cases. On the one hand, if the missing data refer to December, we consider for this month the observation of November of the same year. Similarly, if the missing data refer to January, we consider for this month the observation of February of the same year. On the other hand, if the missing data do not refer to December or January, i.e., if the months before and after the missing observations belong to the same year, a linear interpolation is used.

Finally, it is important to note that credibility indexes need to be interpreted with caution for four countries (the Czech Republic, South Africa, South Korea, and Thailand). Indeed, as we report in table 1, they used or have used a core measure of inflation for their operational target, while inflation expectations published by Consensus Economics address headline inflation.

4.2 Overview of the new credibility indicators

Figures 4 to 7 in the appendix represent $CRED_{LLR1}$ and $CRED_{LLR2}$ for each country, as well as expected and actual inflation and inflation targets. We first observe that negative deviations are very rare. $CRED_{LLR2}$ is then very close to $CRED_{LLR1}$. Moreover, we can see a high correlation between actual and expected inflation. The latter rarely overshoots the former. This means that the monetary authorities in the countries we investigate are generally credible. However, episodes of important loss in credibility are possible.

8 For example, Lysiak et al. (2007) find that consumers and commercial bank analysts form their inflation expectations in very different ways in Poland. While consumers rely heavily on the current inflation target, commercial bank analysts closely follow the announced inflation target. According to Lysiak et al. (2007), this difference could be explained by the costs of collecting and processing information.

Table 1: IT adoption dates, target measures and data coverage

Country	Effective IT start	Target measure	Consensus Economics (Data coverage)	Nb. ofobs. (in months)
Brazil	1999M6	Headline Inflation	1990M2 (monthly since 2001M4)	175
Chile	1999M9	Headline Inflation	1993M3 (monthly since 2001M4)	172
Colombia	1999M9	Headline Inflation	1993M3 (monthly since 2001M4)	172
the Czech Rep.	1998M1	Headline Inflation since 01/2002	1995M1 (monthly since 2007M5)	192
Guatemala	2005M1	Headline Inflation	2009M1 (monthly since 2009M1)	60
Hungary	2001M6	Headline Inflation	1990M11 (monthly since 2007M5)	151
Indonesia	2005M7	Headline Inflation	1990M11 (monthly since 1990M11)	102
Israel	1997M6	Headline Inflation	1995M1 (monthly since 1995M1)	199
Mexico	2001M1	Headline Inflation	1990M2 (monthly since 2001M4)	156
Peru	2002M1	Headline Inflation	1993M3 (monthly since 2001M4)	144
the Philippines	2002M1	Headline Inflation	1994M12 (monthly since 1994M12)	144
Poland	1998M10	Headline Inflation	1990M11 (monthly since 2007M5)	183
Romania	2005M8	Headline Inflation	1995M1 (monthly since 2007M5)	101
Slovakia	2005M1	Headline Inflation*	1995M1 (monthly since 2007M5)	48
South Africa	2000M2	Headline Inflation since 01/2009	1993M6 (monthly since 1993M6)	167
South Korea	2001M1	Headline Inflation since 01/2007	1990M1 (monthly since 1990M1)	156
Thailand	2000M5	Core inflation	1990M11 (monthly since 1990M11)	164
Turkey	2006M1	Headline inflation	1995M1 (monthly since 2007M5)	96

** joined the Eurozone in January 2009.*
Source: Roger (2009), Hammond (2012) and Central Banks' website.

Before comparing our new indicators with the existing ones, table 2 summarizes the country-by-country evolution of the $CRED_{LLR1}$ index over different sub-periods[9]. In particular, we consider the first 12 and 24 months following the adoption of inflation targeting to assess the initial credibility of the central bank. We also focus on the period between June 2007 and December 2008, which is characterized by a surge in food and energy prices and then by a subsequent increase of inflation in most emerging economies[10]. Three main conclusions can be drawn.

First, most emerging inflation-targeting countries display a relatively high level of central bank credibility over the full period we consider (IT start-2013M12). Indeed, the average value of the $CRED_{LLR1}$ index is equal to 0.89, while the average probability that $CRED_{LLR1}$ exceeds 0.95 is equal to 0.67. South Korea exhibits the highest level of credibility, with an index equal to 0.99 on average. Furthermore, as suggested by the fourth column of the table (IT start - 2009M12), these good results are not driven by the recent low inflation environment in the wave of the Great Recession. Our results are consistent with previous empirical studies showing that the adoption of an inflation-targeting framework in emerging economies has helped to better anchor private-sector inflation expectations (see, e.g., IMF (2008), Davis (2014)) and to reduce their dispersion (Capistran & Ramos-Francia (2010)).

Second, focusing on the one year and the two years following the adoption of inflation targeting (columns 1 and 2 of table 2), it appears that the introduction of this new monetary framework was initially perceived as not very credible (Romania, Turkey), if not non-credible (the Czech Republic, Indonesia) by the private sector. Such an initial lack of credibility could be explained by the fact that these countries did not fully satisfy the macroeconomic and institutional preconditions for adopting inflation targeting, such as central bank independence and transparency, fiscal discipline, or exchange rate flexibility. More importantly, the Turkish experience shows that the initial lack of central bank credibility has led to a loss in inflation control and a self-sustaining loss in credibility. To stop this vicious cycle and to reduce the risk of future overshooting, Turkey decided in June 2008 to revise its target upward. However, as we can see in figure 8, this revision was insufficient to restore the medium-run credibility of the central bank. The private sector considered this revision to be a renouncement of the authorities' commitment to price stability (Habermeier et al. (2009)).

[9] As $CRED_{LLR1}$ and $CRED_{LLR2}$ are very close to each other, we only present the characteristics of the former and we will only focus on $CRED_{LLR1}$ in the following sections.

[10] According to Habermeier, Otker-Robe, Jacome, Giustiniani, Ishi, Vavra, Kisinbay & Vazquez (2009), this inflationary episode was the first significant test for the credibility of the inflation targeting regimes in emerging countries.

Finally, it appears that the food and energy price shocks in the second half of 2007 and 2008 did not abruptly destroy the monetary policy credibility of the emerging inflation targeting countries. Indeed, while figures 4 to 8 in the appendix show that most countries overshot their targets during this period, the third column of the table 2 (2007 M6-2008 M12) does not highlight a sharp decrease of the $CRED_{LLR1}$ index, except for Hungary. Of course, the size of the increase in inflation expectations and the evolution of the credibility index depend on the severity of the inflation shocks. Nonetheless, some countries (Chile, Israel, Mexico, South Africa, and South Korea) succeeded in containing inflation expectations notwithstanding a subsequent increase in actual inflation, above the targets. This demonstrates how important well-established past credibility is to addressing adverse supply shocks and to limiting second round effects on output. The lower the credibility is, the stronger the tightening of monetary policy should be (Alichi & Al. (2009), Neuenkirch & Tillmann (2014)).

Table 2: Some characteristics of the $CRED_{LLR1}$ index

	First 12 months	First 24 months	Mean (07M6-08M12)	Mean (IT start - 09M12)	Mean (overall period)	St. Dev. (overall period)	Prob[LLRI > 0.95] (overall period)	Prob[LLRI < 0.5] (overall period)	Rank (overall period)
Brazil	1.00	1.00	1.00	0.91	0.94	0.05	0.87	0.06	9c
Chile	1.00	1.00	0.83	0.97	0.98	0.01	0.91	0.01	2
Colombia	0.76	0.88	0.90	0.96	0.98	0.01	0.88	0.00	4
the Czech Rep.	0.21	0.59	0.78	0.88	0.92	0.04	0.71	0.05	11
Guatemala	-	-	-	-	0.98	0.00	0.82	0.00	-
Hungary	1.00	0.99	0.25	0.66	0.64	0.09	0.19	0.32	17
Indonesia	0.41	0.69	0.55	0.68	0.80	0.09	0.52	0.17	13
Israel	1.00	0.87	1.00	0.97	0.97	0.01	0.87	0.00	5
Mexico	0.91	0.96	0.97	0.98	0.98	0.00	0.85	0.00	3
Peru	1.00	1.00	0.76	0.95	0.97	0.01	0.88	0.03	6
the Philippines	0.92	0.89	0.69	0.75	0.83	0.06	0.50	0.16	12
Poland	0.99	0.82	0.97	0.95	0.96	0.01	0.81	0.00	7
Romania	0.75	0.87	0.60	0.76	0.77	0.05	0.32	0.13	14
Slovakia	0.99	0.87	0.47	0.71	0.71	0.06	0.15	0.19	16
South Africa	1.00	1.00	0.66	0.90	0.93	0.03	0.73	0.04	10
South Korea	1.00	1.00	0.97	0.99	0.99	0.00	0.93	0.00	1
Thailand	1.00	1.00	0.81	0.95	0.95	0.01	0.68	0.02	8
Turkey	0.66	0.56	0.38	0.59	0.72	0.10	0.46	0.27	15
Mean	0.86	0.88	0.74	0.86	0.89	0.03	0.67	0.08	
Median	0.99	0.89	0.78	0.91	0.94	0.02	0.77	0.04	

Note: Guatemala is excluded from the ranking because of the short sample of available data.

4.3 Comparison with the existing indicators

Figures 4 to 7 in the appendix allow for comparing $CRED_{LLR1}$ and $C\ RED_{LLR2}$ to the two existing credibility indicators that constitute a reference so far, namely $CRED_{KM}$ and $CRED_{DMGS}$. Immediately, it appears that the variation amplitudes of $CRED_{LLR1}$ and $CRED_{LLR2}$ are higher than those of the existing indicators inside of the [0-1] interval. This is justifiable. Consider for instance the case of Brazil in 2003-2004, when the monetary authorities entirely lost their credibility according to $CRED_{LLR1}$ and $CRED_{LLR2}$. In 2003, the agents unequivocally (and rightly) expected that the central bank would not meet its commitment. The deviations were important: the expected (actual) inflation rate reached 11% (17%), while the target ceiling was 6.5%. Such a situation encourages agents to ignore the target when negotiating their salary or updating their prices, all the more because of a need to catch up wages and prices, as the initial surge in prices was not covered in the previous contracts. Thus, the context of Brazil in 2003-2004, by definition and given the size of the deviations, can reasonably be considered to be a complete loss in credibility. On the contrary, $CRED_{KM}$ and $CRED_{DMGS}$ did not fall below 0.52 and 0.60, respectively. Given the [0,1] range, the plausibility of their message in terms of credibility is questionable.

The Romanian case also supports our indicators. Indeed, while the monetary authorities failed most of the time to meet the target, and consequently private expectations were often above the target ceiling, $CRED_{KM}$ and $CRED_{DMGS}$ always remained higher than 0.8. However, $CRED_{LLR1}$ and $CRED_{LLR2}$ appropriately address the deviations, falling for instance to 0.40 in 2006, 0.22 in 2008, 0.40 in 2011 and 0.45 in 2013. They plausibly suggest that the Romanian central bank had suffered from loss in credibility.

Turkey also offers an interesting comparison. From 2006 to 2008, the actual and expected inflation rates always exceeded the target ceiling, by up to 4 and 6.5 percentage points, respectively. In these conditions, it is very hard to believe that the Turkish central bank was credible over this period. However, $CRED_{KM}$ and $CRED_{DMGS}$ remained close to 0.75 on average. Once again, to the contrary, $CRED_{LLR1}$ and $CRED_{LLR2}$ duly signal a significant loss in credibility, with values considerably lower than 0.5. Interestingly, while our indicators highlight a full loss in credibility at the end of 2008, Turkey decided to raise the target range, a maneuver to restore credibility (see, e.g., Habermeier et al. (2009)).

Similarly, the existing indicators do not appropriately address the (sometimes huge) deviations of inflation expectations from the target that typically occurred in South Africa, Indonesia, the Philippines and Hungary, contrary to $CRED_{LLR1}$ and $CRED_{LLR2}$.

5 The impact of credibility on the volatility of monetary policy instrument

We now investigate the extent to which central bank credibility, as measured by our new indicator, influences the volatility of the key instrument of monetary policy, namely the short-term interest rate. This is an important issue, as a credible central bank is more likely to anchor inflation expectations to its target. In such a case, the central banker does not have to move his key instrument too much to influence the yield curve in the desired direction. At the extreme, speeches are enough. On the contrary, non-credibility is penalizing in that it implies more volatility of the interest rate, while the variance of the interest rate theoretically enters the micro-founded welfare-based loss function of central banks[11]. Furthermore, the volatility of the monetary policy instrument increases macroeconomic uncertainty and (financial) instability. Thus, we want to test the following hypothesis: a higher (lower) credibility contributes to a lower (higher) volatility of the interest rate.

A similar issue has been addressed by De Mendonca & de Guimarâes e Souza (2009) in the case of Brazil. However, they do not explicitly assess the relationship between credibility and interest rate volatility, as they regress the first-difference of the interest rate on the variation of their credibility index by using Ordinary Least Squares (OLS) estimates. We consider a General Auto-Regressive Conditional Heteroscedastic (GARCH) approach to be more adequate for analyzing the volatility of any variable. We will use such a model to test whether our index of credibility significantly influences the conditional variance of interest rates.

[11] See the demonstration of Woodford (2003, Chap. 6).

First, it is common to consider interest rate rules, such that the short-term interest rate responds to the deviations of the inflation rate towards its target, with a gradual adjustment (see, e.g., Clarida, Gali & Gertler (1998)). Consistent with this, the mean equation of our GARCH specification is a second order autoregressive process[12] augmented with the inflation rate and with a constant that is supposed to represent both the inflation target and the long-run equilibrium interest rate:

$$i_t = c + \rho_1 i_{t-1} + \rho_2 i_{t-2} + \phi \pi_t + \varepsilon_t \tag{11}$$

ε_t represents the innovations of the short-term interest rate (free of inflationary shocks) at time t with a zero mean and time-varying variance h_t. More precisely, we suppose that $\varepsilon_t = z_t \sqrt{h_t}$, with z_t representing a standardized white noise residual.

The time-varying conditional variance of the interest rate is supposed to follow an Exponential General Auto-Regressive Conditional Heteroscedastic (EGARCH) process, augmented with the lagged *CRED-LLR1* indicator as an additional determinant. Its general representation is given by:

$$log\,(h_t) = \alpha_0 + \sum_{i=1}^{q} \alpha_i g(z_{t-i}) + \sum_{i=1}^{p} \beta_i log(h_{t-i}) + \omega CRED\_LLR1_{t-1} \tag{12}$$

with $g\,(z_{t-i}) = \theta z_{t-i} + \gamma\,(|z_{t-i}| - E|z_{t-i}|)$, where $E|z_{t-i}|$ is conditional to a given density function. While estimating a GARCH(p,q) model requires the parameters a_i and $в$ to be positive (because variance cannot be negative), the EGARCH(p,q) model is expressed in terms of the log of h_t. Thus the conditional variance will always be positive whatever the sign of the parameters (Nelson (1991)). This is important in our specific case because *CRED-LLR1* is expected to have a negative influence on the conditional variance of the interest rate (namely, *ш* is expected to be negative).

The first column of table 5 in the appendix reports the results of excess kurtosis tests for the interest rate data series. The null hypothesis of normality is only rejected for Colombia, the Czech Republic and Mexico. For these three countries, a Student's *t* distribution with a degree of freedom *v* (to be estimated) is then preferred to a normal one, as is usual in the case of leptokurtic distribution.

Table 5 in the appendix also reports the results of no ARCH effect tests. Such a test requires serially uncorrelated e_t. However, the usual *Q* tests of no serial correlation rely on an assumption of conditional homoscedasticity. So we used the "robust" *Q* test suggested by West & Cho (1995). As indicated in the fourth column, the null hypothesis of an absence of serial correlation is not rejected at the usual risk levels for every country, even if we have some doubt about Peru. Finally, the hypothesis of no ARCH effect (for lags = 2, 4 and 6 months) is clearly rejected for most of the countries, except for Hungary, Israel, Slovakia, Thailand and Turkey. For the other countries, the interest rate data series exhibit types of large residual clustering that is consistent with a GARCH specification.

Tables 3 and 4 report the results of the estimations of the EGARCH(1,1)-X models[13].

[12] Considering an autoregressive process higher than the first order allows us to remove the serial correlation of residuals as well.

[13] Hungary, Israel, Slovakia, Thailand and Turkey are excluded because of the absence of ARCH effects, while Guatemala is not considered because of the lack of short-term interest rate data series.

Table 3: EGARCH-X estimates (1/2)

	Brazil	Chile	Colombia	the Czech R.	Indonesia	Mexico
MEAN EQUATION						
constant	*0.063*	*0.107****	*0.017****	*-0.022**	*0.324****	*-0.007*
	(0.067)	*(0.025)*	*(0.006)*	*(0.012)*	*(0.012)*	*(0.031)*
i_{t-1}	*1.802****	*1.559****	*1.414****	*1.267****	*0.896****	*1.309****
	(0.016)	*(0.067)*	*(0.001)*	*(0.064)*	*(0.019)*	*(0.078)*
i_{t-2}	*-0.815****	*-0.589****	*-0.427****	*-0.286****	*-0.069****	*-0.313****
	(0.018)	*(0.005)*	*(0.001)*	*(0.062)*	*(0.017)*	*(0.078)*
π_t	*0.019***	*0.017****	*0.009****	*0.028****	*0.097****	*0.006*
	(0.009)	*(0.005)*	*(0.001)*	*(0.006)*	*(0.003)*	*(0.008)*
VARIANCE EQUATION						
constant	*-0.632***	*0.354*	*0.770****	*0.028*	*-2.304****	*-0.982**
	(0.278)	*(0.396)*	*(0.001)*	*(0.432)*	*(0.264)*	*(0.524)*
$g\ (z_t\text{-}i)$	*0.497****	*0.661****	*-0.291****	*2.061*	*2.287****	*0.503***
	(0.114)	*(0.104)*	*(0.001)*	*(1.394)*	*(0.203)*	*(0.239)*
$h_t\text{-}1$	*0.643****	*0.949****	*0.942****	*0.897****	*0.865****	*1.011****
	(0.013)	*(0.015)*	*(0.001)*	*(0.039)*	*(0.050)*	*(0.009)*
$CRED\_LLR1_{t-1}$	*-0.729****	*-1.043****	*-0.788****	*-0.564**	*0.262*	*0.607*
	(0.260)	*(0.385)*	*(0.001)*	*(0.313)*	*(0.375)*	*(0.525)*
Degrees of freedom (a)	-	-	*2.92*	*2.04*	-	*2.42*
GARCH LB test (b)	*0.078*	*0.586*	*0.035*	*0.213*	*0.526*	*0.999*
GARCH McLL test (c)	*0.994*	*0.774*	*0.750*	*0.643*	*0.318*	*0.999*
Number of observations	*173*	*167*	*170*	*163*	*100*	*132*

*Notes: Std. errors are in parentheses. *, **, and *** denote signiflcance at the 10%, 5% and 1% level, respectively.*
(a) Estimation of the number of degrees of freedom v (in case of Student-t distribution).
(b) P-Value of the Ljung-Box no serial correlation test on the standardized residuals £t/\/(ht).
(c) P-Value of the McLeod-Li no serial correlation test on the squared standardized residuals e2/ht.

Focusing on the variance equation, the nullity of α_1 and β_1 is rejected for every country, except for the nullity of α_1 for the Czech Republic and South Korea. So, the current conditional variance of the interest rate is significantly explained by past innovations contained both in *g(.)* and in the past conditional variance $h_{t-1.}$ This confirms the existence of ARCH effects and supports our econometric approach. Moreover, according to the test suggested by McLeod & Li (1983), the null hypothesis of no serial correlation for the squared standardized residuals is never rejected at the usual risk levels. This suggests that the variance equation is correctly specified with the orders q =1 and p =1 chosen for the EGARCH.

In the same way, a second-order autoregressive process is found to be appropriate for the mean equation. Indeed, according to the usual Ljung & Box test, the null hypothesis of no serial correlation test for the standardized residuals is never rejected at the usual risk levels. Finally, the fact that the estimated degrees of freedom are low (close to but higher than 2) for Colombia, the Czech Republic and Mexico validates *ex post* the choice of a Student's *t* distribution.

Despite the very important informational content of the past conditional variance h_{t-1}, we find that the estimated coefficient associated with $CRED_{LLR1}$ is always statistically significant, except for Indonesia and Mexico, and negative. This confirms that central bank credibility decreases the volatility of the key instrument of monetary policy. In that sense, credibility improves the efficiency of monetary policy, notably through the expectations channel.

Finally, one can argue that central bank credibility evolves according to a sluggish process, in that it can rarely be suddenly increased or annihilated (see, e.g., Blinder (2000)). In this respect, as robustness checks, we have substituted the 6-month moving average of $CRED_{LLR1}$ for its one-lagged value in the variance equation of the EGARCH models. The corresponding results are reported in tables 6 and 7 in the appendix. Overall, they are robust to this alternative specification, even if the estimated coefficient associated with the credibility becomes insignificant for Peru and Poland.

Table 4: EGARCH-X estimates (2/2)

	Peru	the Philippines	Poland	Romania	South Africa	South Korea
MEAN EQUATION						
constant	0.286***	-0.044	-0.069	0.077	0.014***	0.032
	(0.003)	(0.030)	(0.059)	(0.148)	(0.001)	(0.028)
i_{t-1}	1 727***	1.305***	0.845***	1.326***	1.573***	1.542***
	(0.001)	(0.001)	(0.009)	(0.072)	(0.001)	(0.004)
i_{t-2}	-0.783***	-0.303***	0.116***	-0.345***	-0.584***	-0.549***
	(0.001)	(0.002)	(0.015)	(0.081)	(0.001)	(0.010)
π_t	-0.019***	0.004	0.077***	0.017	0.013***	-0.001
	(0.001)	(0.007)	(0.007)	(0.019)	(0.001)	(0.006)
VARIANCE EQUATION						
constant	-2.253***	-0.242	-0.097*	-0.255	-0.274	7.659***
	(0.077)	(0.173)	(0.051)	(0.381)	(0.285)	(1.752)
$g\ (z_t\text{-}i)$	1.517***	0.373***	0.328***	1.062***	0.546***	-0.003
	(0.074)	(0.101)	(0.106)	(0.213)	(0.109)	(0.097)
$h_t\text{-}1$	0.538***	0.866***	0.959***	0.567***	0.767***	0.735***
	(0.019)	(0.051)	(0.016)	(0.123)	(0.080)	(0.063)
CRED1- LLR_{t-1}	-0.508***	-0.617***	-0.262***	-1.142**	-0.985***	-9.032***
	(0.078)	(0.202)	(0.049)	(0.573)	(0.302)	(1.997)
Degrees of freedom (a)	-	-	-	-	-	-
GARCH LB test (b)	0.227	0.501	0.119	0.783	0.491	0.688
GARCH McLL test (c)	0.996	0.321	0.682	0.982	0.184	0.557
Number of observations	131	142	181	99	165	154

*Notes: Std. errors are in parentheses. *, **, and *** denote significance at the 10%, 5% and 1% level, respectively.*
(a) Estimation of the number of degrees of freedom v (in case of Student-t distribution).
(b) P-Value of the Ljung-Box no serial correlation test on the standardized residuals Et/^/(ht).
(c) P-Value of the McLeod-Li no serial correlation test on the squared standardized residuals e2/ht.

6 Concluding remarks

The aim of this article was to provide a simple time-varying metric of central bank credibility. To this end, we suggest a measure of credibility based on the gap between private sector inflation expectations and the inflation target. In contrast to the existing measures, our index introduces two major innovations. First, it is an asymmetric measure of credibility that is based on a linear-exponential (LINEX) function. Indeed, one can expect that, in practice, negative deviations of inflation expectations from the target are less likely to indicate a loss in credibility than positive deviations. Second, contrary to the main contributions to date, our measure does not impose any *ad hoc* threshold above which credibility is considered to be null.

We then compute our index for all emerging inflation-targeting countries and compare it to the existing indicators. Our findings suggest a relatively high level of central bank credibility in these countries over the inflation targeting period. Nonetheless, we observe that monetary policy was not necessarily perceived to be very credible in the immediate wake of inflation targeting adoption, in particular in the Czech Republic, Indonesia, Romania, and Turkey. More importantly, we show that our measure is more suited to assessing the monetary experiences of these economies than the existing ones. In particular, our index is better able to discriminate between the periods of low *versus* high credibility in a context of rather low inflation targets.

Finally, we empirically investigate the linkage between central bank credibility (measured by our index) and short-term interest rate volatility. An EGARCH model is used to this end. Our results confirm that the level of credibility negatively impacts the variance of the interest rate in a large number of countries. This suggests that a credible central bank does not need to frequently change its key instrument to reach the inflation target. Credibility is then expected to improve monetary policy transmission efficiency, particularly through the expectations channel. In terms of policy implications, this means that candidates for an inflation-targeting framework need to make institutional reforms that will ensure an initial high level of credibility. Otherwise, an initial weak credibility could lead to higher and self-sustaining volatility in interest rates (as indicated by our GARCH experiments), which in turn would trigger higher macroeconomic instability.

Against this background, an interesting extension would consist of investigating the economic and institutional factors ensuring a minimum level of credibility. Combining our suggested credibility index with the literature on the preconditions for adopting inflation targeting would be relevant for such research. Indeed, the literature highlights some de- terminants that are likely to play a role, such as the degree of independence of the central bank, the fiscal context, the exchange rate regime, and the quality of institutions. Revealing the deep factors that establish initial credibility is very important for those emerging countries that are candidates for the inflation-targeting framework.

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Appendix

Figure 4: Target range, expected in ation and credibility indicators (1/5)

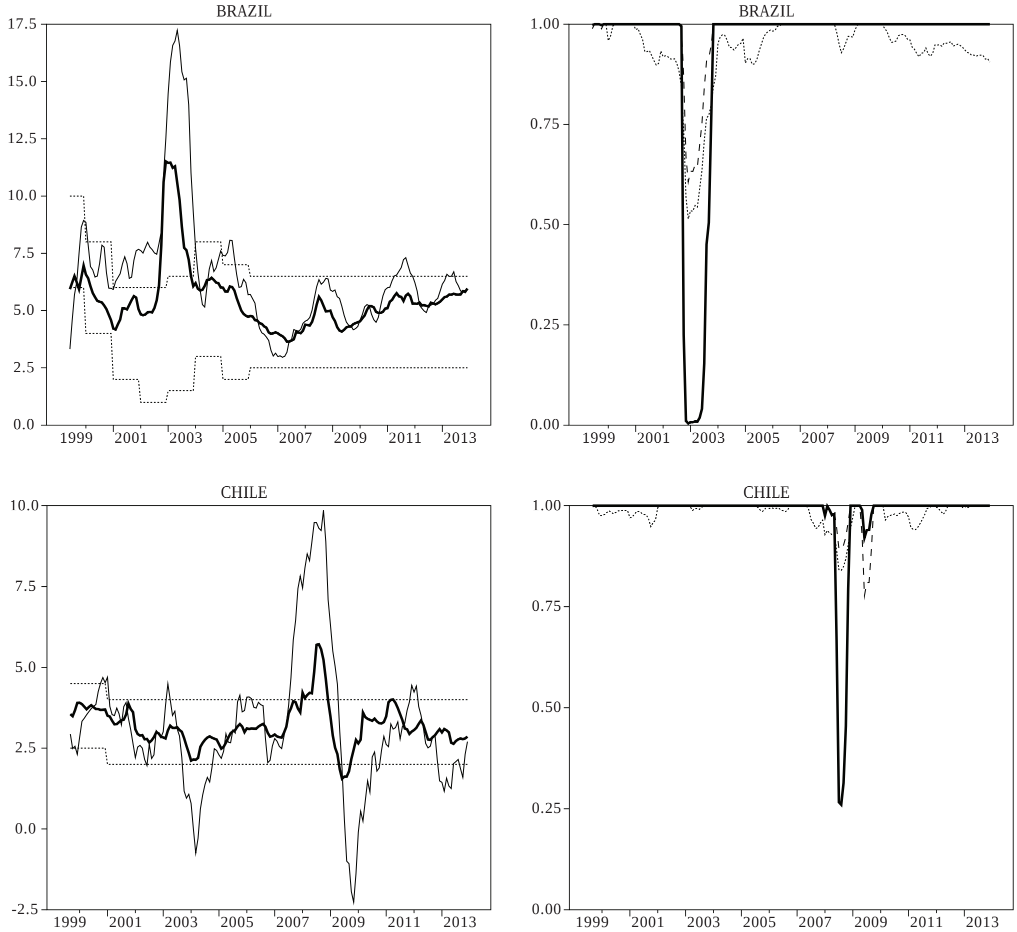

Figure 5: Target range, expected in ation and credibility indicators (2/5)

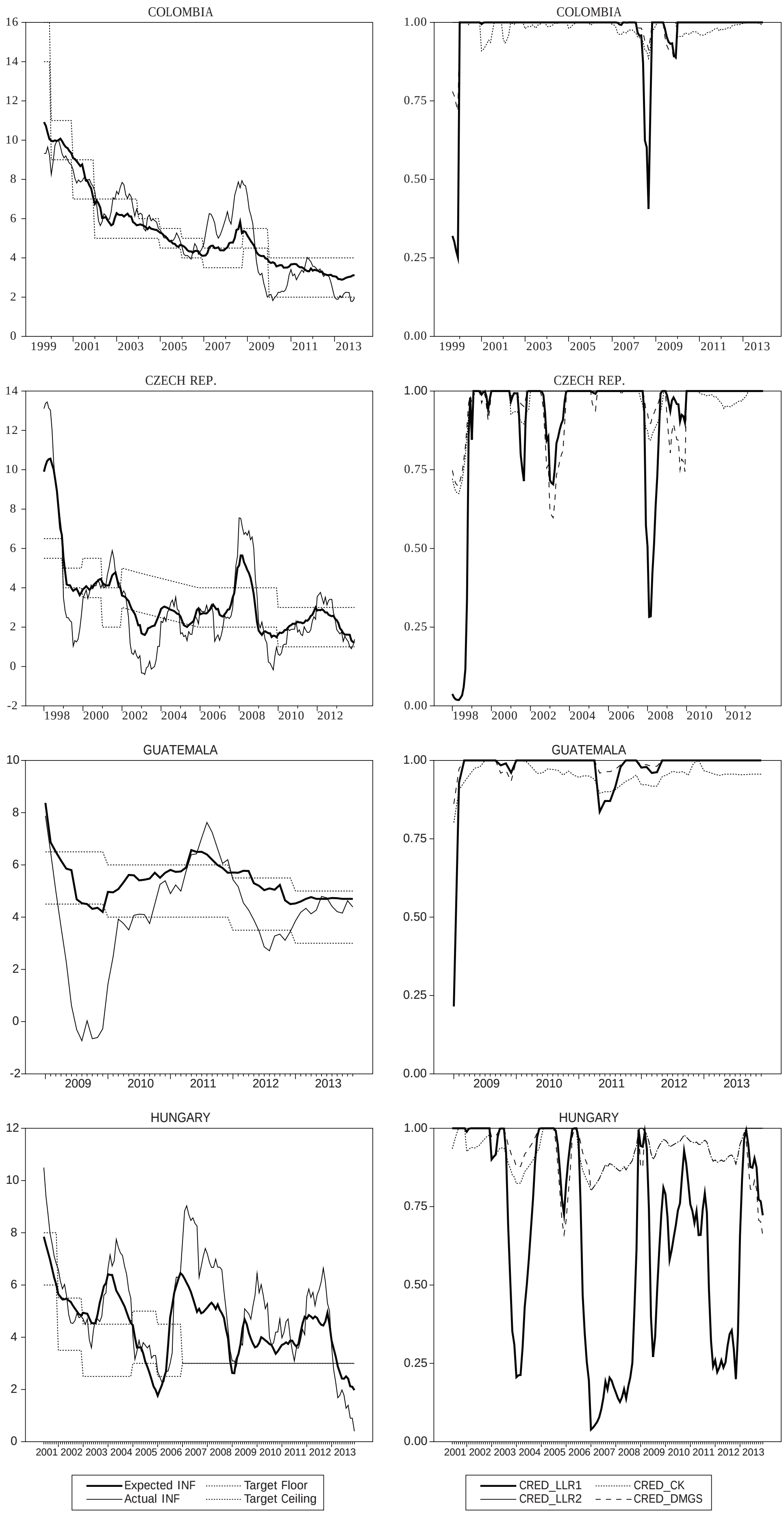

Figure 6: Target range, expected in ation and credibility indicators (3/5)

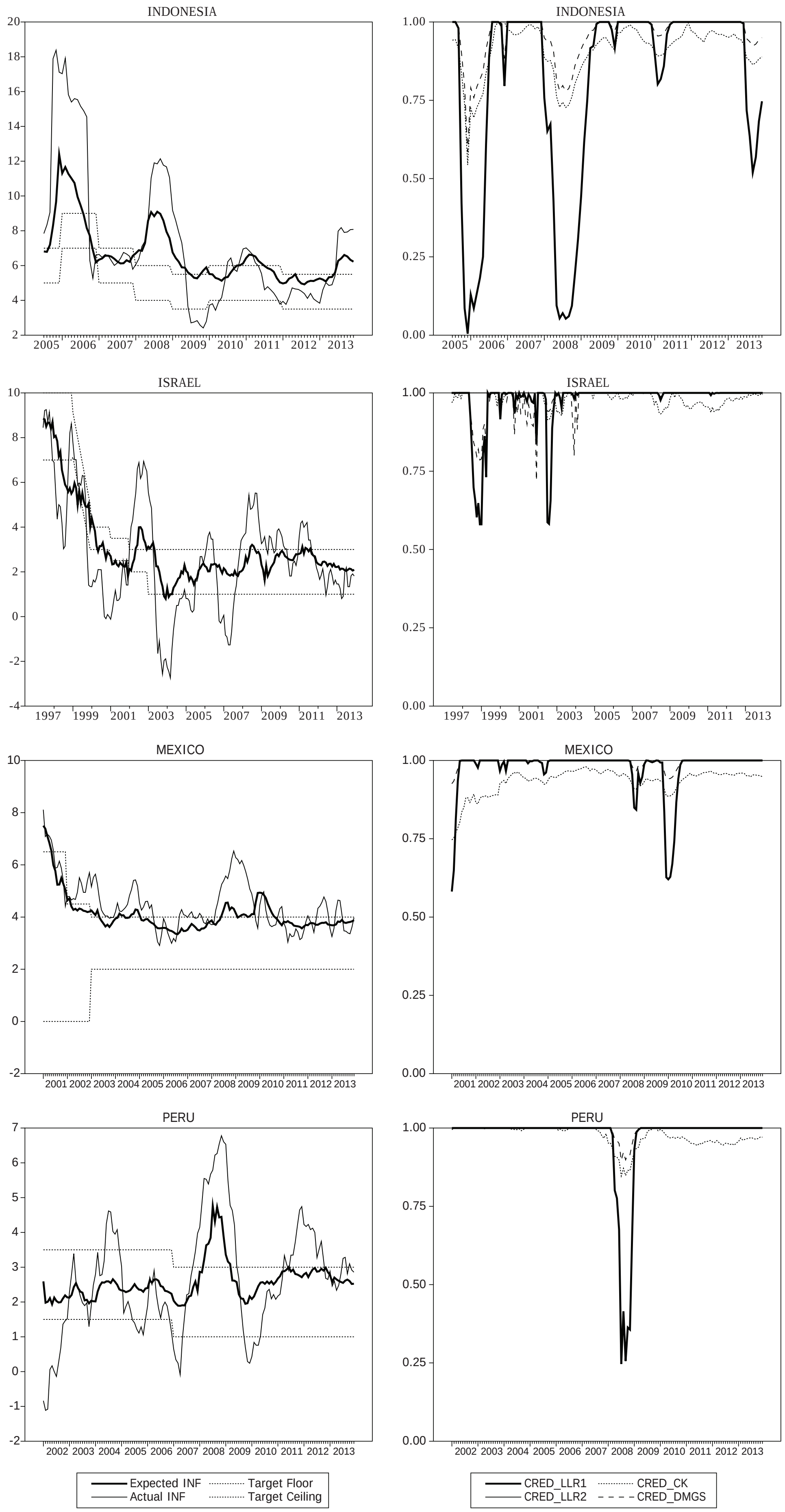

Figure 7: Target range, expected in ation and credibility indicators (4/5)

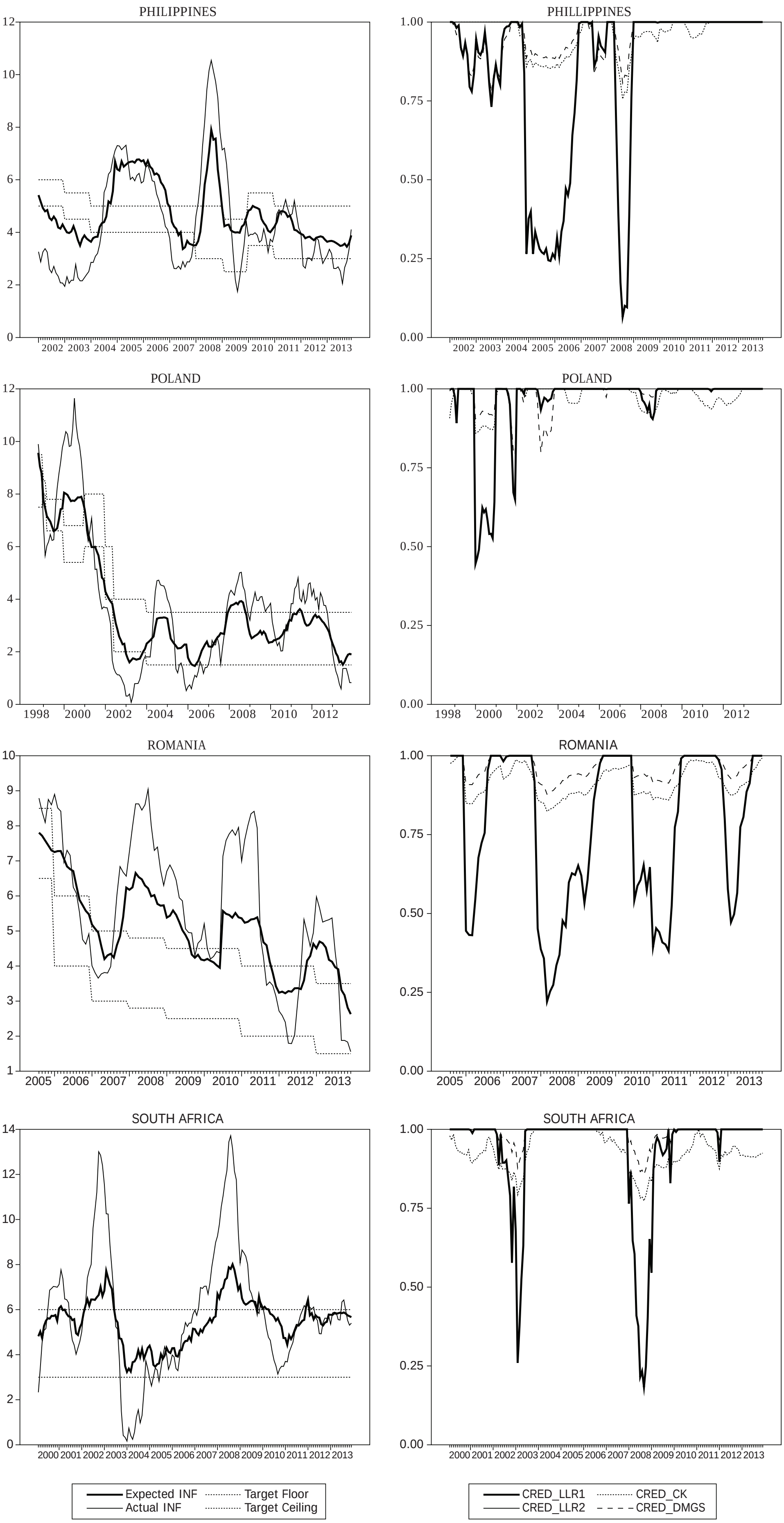

Figure 8: Target range, expected in ation and credibility indicators (5/5)

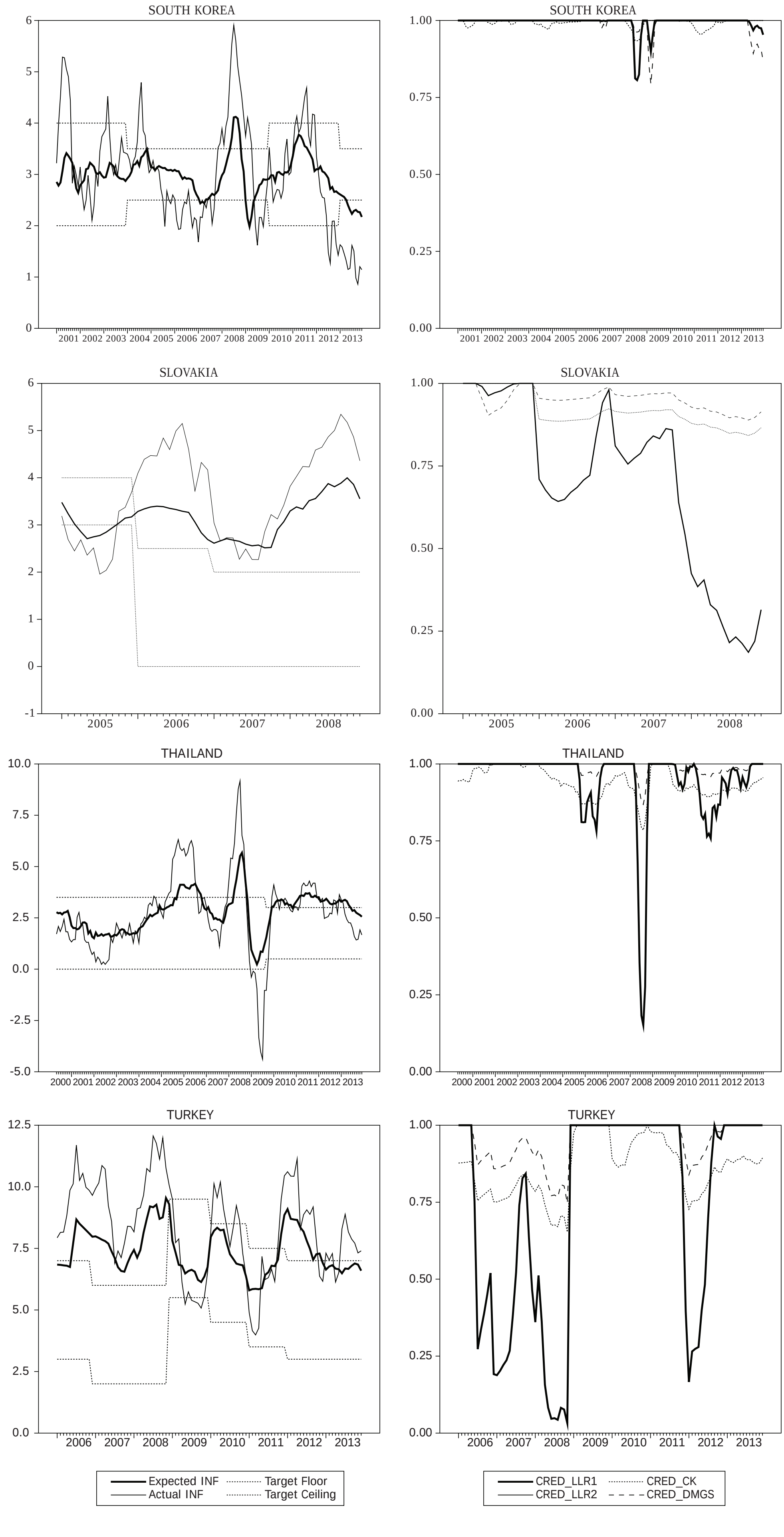

Table 5: Properties of the interest rate data series and tests on the mean equation residuals

Country	Kurtosis excess on interest rate data series (a)	No serial correlation test on residuals εt (b)	No ARCH Effect test on residuals εt (c)		
			lags = 2	lags = 4	lags = 6
Brazil	-0.52	0.764	0.022	0.023	0.056
Chile	0.42	0.591	0.000	0.000	0.000
Colombia	0.88	0.035	0.000	0.017	0.000
the Czech Rep.	5.51	0.918	0.000	0.000	0.000
Hungary	-0.34	0.168	0.951	0.982	0.000
Indonesia	-0.47	0.708	0.000	0.000	0.000
Israel	-0.56	0.846	0.137	0.403	0.096
Mexico	4.32	0.605	0.006	0.000	0.000
Peru	-0.05	0.001	0.001	0.008	0.035
the Philippines	-0.89	0.547	0.056	0.001	0.004
Poland	0.13	0.563	0.000	0.000	0.000
Romania	0.33	0.938	0.025	0.039	0.127
Slovakia	-0.96	0.101	0.233	0.248	0.398
South Africa	-0.95	0.672	0.111	0.002	0.014
South Korea	-1.01	0.965	0.005	0.005	0.021
Thailand	-0.32	0.100	0.115	0.256	0.052
Turkey	-1.60	0.176	0.305	0.596	0.912

*(a) * means rejection of the Normality hypothesis at the 5% level (leptokurtic distribution).*
(b) P-value of the West & Cho (1995) test on the residuals εt of the mean equation.
(c) P-value of the ARCH test consisting in regressing the square residuals series on its own lags. Under the null, the corresponding R^2 is equal to zero.

Table 6: EGARCH-X estimates with the 6-month moving average of $CRED_{LLR1}$ (1/2)

	Brazil	Chile	Colombia	the Czech R.	Indonesia	Mexico
MEAN EQUATION						
constant	0.040*	-0.054***	0.003	-0.020	0.272**	-0.010
	(0.021)	(0.005)	(0.029)	(0.014)	(0.111)	(0.029)
it−1	1.812***	1.558***	1.512***	1.274***	1.138***	1.307***
	(0.001)	(0.001)	(0.059)	(0.073)	(0.020)	(0.004)
it−2	-0.821***	-0.557***	-0.520***	-0.293***	-0.227***	-0.311***
	(0.001)	(0.002)	(0.061)	(0.071)	(0.023)	(0.001)
πt	0.005	0.018***	0.010	0.027***	0.027	0.006
	(0.004)	(0.003)	(0.011)	(0.007)	(0.023)	(0.006)
VARIANCE EQUATION						
constant	-0.699***	0.559***	1.499***	-0.239	-1.026***	-1.337**
	(0.100)	(0.017)	(0.126)	(0.367)	(0.331)	(0.598)
g (zt−1)	0.837***	0.602***	1.195***	0.635**	1.036***	1.130
	(0.149)	(0.032)	(0.136)	(0.264)	(0.383)	(0.862)
ht−1	0.156***	0.943***	0.964***	0.863***	0.951***	0.998***
	(0.036)	(0.004)	(0.021)	(0.054)	(0.063)	(0.013)
CRED LLR1 MA(6)	-2.378***	-1.207***	-1.604***	-0.729*	0.129	0.976
	(0.093)	(0.017)	(0.140)	(0.451)	(0.202)	(0.622)
Degrees of freedom (a)	-	-	2.01	2.88	-	2.09
GARCH LB test (b)	0.060	0.240	0.004	0.645	0.986	0.556
GARCH McLL test (c)	0.805	0.742	0.980	0.628	0.984	0.154
Number of observations	170	167	167	160	97	130

*Notes: Std. errors are in parentheses. *, **, and *** denote significance at the 10%, 5% and 1% level, respectively.*
(a) Estimation of the number of degrees of freedom υ (in case of Student-t distribution).
(b) P-Value of the Ljung-Box no serial correlation test on the standardized residuals $\varepsilon_t\sqrt(h_t)$.
(c) P-Value of the McLeod-Li no serial correlation test on the squared standardized residuals ε_t^2/h_t.

Table 7: EGARCH-X estimates with the 6-month moving average of CREDLLR1 (2/2)

	Peru	the Philippines	Poland	Romania	South Africa	South Korea
MEAN EQUATION *constant*	0.131***	-0.048	-0.048	0.180***	-0.036*	0.053***
	(0.026)	(0.033)	(0.077)	(0.038)	(0.021)	(0.001)
it−1	1.757***	1.253***	0.834***	1.328***	1.499***	1.618***
	(0.051)	(0.003)	(0.037)	(0.005)	(0.001)	(0.001)
it−2	-0.787***	-0.250***	0.124***	-0.350***	-0.507***	-0.625***
	(0.048)	(0.004)	(0.043)	(0.005)	0.002	(0.001)
πt	-0.009	0.005	0.074***	0.001	0.015***	-0.007***
	(0.009)	(0.007)	(0.010)	(0.008)	(0.004)	(0.000)
VARIANCE EQUATION						
constant	-6.256***	-0.288	-0.045	-0.057	-0.105	-3.497***
	(1.628)	(0.232)	(0.368)	(0.094)	(0.144)	(0.075)
g (zt−1)	0.915***	0.366***	0.335**	1.090***	0.533***	0.502***
	(0.083)	(0.139)	(0.136)	(0.128)	(0.136)	(0.109)
ht−1	-0.521***	0.825***	0.955***	0.546***	0.630***	-0.566***
	(0.029)	(0.091)	(0.020)	(0.091)	(0.047)	(0.087)
CRED LLR1 MA(6)	0.227	-0.727***	-0.333	-1.449***	-1.660***	-4.259***
	(1.698)	(0.256)	(0.326)	(0.147)	(0.319)	(0.381)
Degrees of freedom (a)	-	-	-	-	-	-
GARCH LB test (b)	0.426	0.400	0.110	0.698	0.322	0.435
GARCH McLL test (c)	0.999	0.190	0.705	0.954	0.108	0.004
Number of observations	128	139	178	96	162	151

*Notes: Std. errors are in parentheses. *, **, and *** denote significance at the 10%, 5% and 1% level, respectively.*
(a) Estimation of the number of degrees of freedom v (in case of Student-t distribution).
(b) P-Value of the Ljung-Box no serial correlation test on the standardized residuals $\varepsilon_t \sqrt{(h_t)}$.
(c) P-Value of the McLeod-Li no serial correlation test on the squared standardized residuals ε_t^2 / h_t .

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SENSITIVITY OF INFLATION TO DEMAND CONDITIONS IN TURKEY: DETERMINING CPI ITEMS RESPONDING TO OUTPUT GAP AND CREDITS*

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Abstract

This study aims to determine CPI items that are influenced by output gap or credits in order to measure the sensitivity of inflation to demand conditions in Turkey. We employ a disaggregated strategy where we first estimate different Phillips curves for 152 sub-indices of the CPI. Then we identify the goods and services through which the effects of credit and output gap on inflation are observed. Our results reveal that one-third of the CPI significantly responds to output gap, while about one-fourth of the CPI significantly responds to credit use in Turkey. Overall, we conclude that economic policies targeting aggregate demand and credit may influence half of the inflation basket. This study will also contribute to compilation of core inflation measures that would enable policymakers to better track the effects of monetary policy on inflation.

* This work is based on two studies: Özmen and Sankaya (2014) and Atuk, Aysoy, Özmen and Sankaya (2014). The views and opinions presented in this study belong to the authors and do not necessarily represent those of the Central Bank of the Republic of Turkey or its staff.

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1. Introduction

Starting with late 2010, the Central Bank of Turkey (CBRT) expanded the set of policy tools with an aim of incorporating financial stability concerns as well as achieving price stability. In this regard, the conventional inflation targeting regime, where demand and expectation management is the key channel and interest rate is the main instrument, has been enhanced with a special focus on credit developments through the use of required reserves.[1] In this framework, the effects of monetary policy on domestic demand and inflation are observed through two channels: (1) Interest rate and liquidity management, (2) credits. Therefore, the stance of monetary policy regarding demand-management cannot be revealed by ignoring any of the two channels. From this perspective, understanding how inflation is related to interim objectives of economic growth and credit growth rate may provide useful insights for selection of right policy instruments.

This study investigates the sensitivity of Consumer Price Index (CPI) sub-items to economic policies targeting aggregate demand via output gap and credits in Turkey. To this aim, standard Phillips curve equations for each of the 152 subgroups of the COICOP 5-digit CPI are estimated and the goods and services prices that respond to output gap or credits in a statistically and economically significant manner are determined. Afterwards, new price indices are compiled accordingly for items influenced by output gap and credits.

Empirical findings show that around 1/4th of the CPI are sensitive to credits; meanwhile, 1/3rd is sensitive to output gap. Even though the number of items affected by credits is higher than the number of items affected by output gap, their relative share in the basket is smaller. Credit-sensitive items are mostly registered under core goods group (goods excluding energy, food, alcoholic beverages, tobacco products and gold) which is sensitive to capital inflows, exchange rate and credit cycles. On the other hand, prices in the services sector, which is more labor intensive and generally closed for foreign trade, are more sensitive to output gap than credits. Overall, our findings suggest that about one half of the CPI is affected either by the output gap or the credits.

The study also sheds light on the lag structure of the monetary transmission through conventional output gap channel and through credit stabilization policy channel, which has become popular recently under concerns of financial stability. While the transmission of output gap to consumer prices is observed in 2 to 3 quarters, the effect of credits is observed in 4 to 6 quarters. Thus, two different demand-management channels are effective over prices at different horizons. Overall, one may conclude that along with the classical output gap channel, the credit policy is also an effective policy tool for prolonging the effects of monetary policy.

The remaining items of the CPI, those not affected by demand-management policies, co- move with imported inflation as they are highly sensitive to exchange rate and foreign price developments. The pass-through of foreign price and exchange rate shocks to those items is almost instantaneous and mostly completed in 1 to 2 quarters. Nonetheless, it is worth noting that items sensitive to output gap and/or credits are also affected by foreign prices to some extent. Especially the core goods inflation, which is found to be highly sensitive to credits, is closely related with exchange rate and foreign price fluctuations due to high import intensity of the group. Therefore, to extract the direct impact of output gap and credits to price indices, the foreign price effects are also needed to be controlled for.

The results offer several policy implications. First, solely resorting to countercyclical policies may not be sufficient for a sustained disinflation in Turkey since half of the consumer basket is found to have no direct linkage with demand conditions. Second, given the limited potency of cyclical policies in affecting inflation reducing impact range of the transmission of foreign price shocks to domestic prices and eliminating product and labor market rigidities should be placed on top of the reform agenda for achieving sustained price stability.

[1] For more information on the recent monetary policy framework of CBRT, reader may refer to Başçı and Kara (2011), Kara (2012), Alper et al. (2013).

2. Data, Methodology and Identification Strategy

The most generally used theoretic framework to associate inflation and demand conditions is the Phillips curve relation. In macro models, the hybrid Phillips curve (PC), which encompasses forward and backward-looking pricing behavior and the evolution of real marginal costs, is generally used[2]. If real marginal costs commove with business cycle (pro- cyclicality of marginal costs) then, output gap can be a good indicator in explaining inflation. Therefore, in the generic New Keynesian PC (NKPC), the output gap will be used as the main indicator of demand pressure in the economy. Yet, output gap is only one of the proxies for business cycles which are not observable directly. Lucas (1977) defined business cycles as an interaction variable incorporating the co-movement of various macro indicators, including but not limited to growth, unemployment, credits.

To our best knowledge, there is no theoretic background relating credits to inflation other than through demand and cost pressure cannels. In other words, credits are not deemed to alter inflation per se. There are several studies claiming that credits may serve the purpose of the output gap in NKPC as a measure of cost measure.[3] In addition, in the literature it is also argued that credit cycles may be leading business cycles.[4] Hence, credits are considered as another proxy of the business cycles rather than a variable directly related. Therefore, in this study we estimate a reduced form of PC with alternating "demand measures" including output gap and credit growth.

In this perspective, we estimate PC equations for 152 sub-indices of the CPI (at COICOP 5-digit level) which include either the output gap or credit measures to proxy demand conditions. The credit measures are defined as the quarterly share of credits to GDP for three different types of credits: consumer, business and total credits. The specification of the PC is as follows: [5]

$$\pi_t = c + \alpha\pi_{t-1} + \beta(GAP\ or\ CREDIT)_{t-i} + \sum_{k=0}^{K} \gamma_k . PMTL_{t-k} + \sum_{j=0}^{J} \delta_j . MinWage_{t-j} + \varepsilon_t$$

$$i = 0, \dots, 6.$$

In this specification, n is the quarterly inflation; gap is the output gap; credit is the credits; *PMTL* and *MinWage* are the control variables referring to import prices in domestic currency and net minimum wage respectively; and finally e is the error term. The identification strategy builds on the detection of sub-indices where the β coefficient is statistically significant and positive, suggesting that the price index is sensitive to output gap or credits depending on the specification. The data definitions are reported in Table 1:

2 See Gali and Gertler (1999).

3 Ravenna and Walsh (2006), Waters (2013).

4 Kiyotaki (1998), Lown and Morgan (2006), Gilchrist and Zakrajsek (2012).

5 Similar analysis has been conducted by Froehling and Loomatzsch (2011) for EU countries and by Halka and Kotlowski (2013) for Poland.

Table 1: Data Definitions

Data	Description	Source
Inflation	Quarterly percent change of price indices	COICOP 5-digit CPI (2003=100) sub-indices, Turkstat
Credits	Quarterly change in credit stock as a percentage of the quarterly GDP	Consumer, business and total credits, CBRT
Output gap	Percent deviation of the GDP from its potential level	Alp, Öğünç and Sarıkaya (2012)
Import prices	Quarterly percent change of the Turkish lira denominated import price index	Import Unit Value Index (in TL) (2010=100), Turkstat
Wages	Quarterly percent change of the minimum wage	Net minimum wage, Ministry of Labor and Social Security

Notes: Consumer credits are excluding housing. Business credits and otal credits are adjusted for exchange rate effects.

The quarterly data covers 2004Q1-2014Q1 period. The generic equation includes import prices in Turkish lira simultaneously and with 4 lags; net minimum wage simultaneously and with one lag and the lagged inflation. Using this base specification, first, simultaneous value and 6 lags of the output gap enter the model. Therefore, for each price index, we estimate 7 different PCs with output gap as the demand measure. Likewise, using simultaneous value and 6 lags of credits for three different credit types, we estimate 21 PCs with credits as the demand measure for each sub-index. Overall, 28 different PC equations are estimated for each of 152 sub price indices (4256 equations in total).[6] Once the models are estimated, the identification of sub-indices sensitive to output gap or credits is achieved as follows: If, for instance, the coefficient of the output gap is statistically and economically significant in any of the 7 equations for an index, we denote that sub price index as being sensitive to output gap. Finally, once the items are determined, then, aggregate price series (i.e. CPI sensitive to output gap) are calculated.[7]

3. Empirical Findings

The findings of the estimation results regarding output gap/credit inflation relation can be summarized under several headings. First of all, out of 152 items of the CPI, 47 are found to be sensitive to output gap, meanwhile 60 items are influenced by credits. However, considering the relative item weights, items that are sensitive to output gap constitute 1/3rd of the basket. Items sensitive to credits correspond to about 1/4th of the basket only, although this group includes a higher number of items in comparison to items sensitive to output gap. The breakdown of the number of items with regards to major CPI groups is presented in Table 2.

Table 2: CPI Sub-Indices Sensitive to Output Gap (Gap) and Credits (Credit)

	Number of items			*Share in CPI (%)*		
	Gap	Credit	Gap or Credit	Gap	Credit	Gap or Credit
Core goods	15	34	42	6.7	16.2	20.1
Services	24	14	31	20.6	4.1	21.4
Food	7	12	15	5.4	5.2	9.2
Energy	1	0	1	1.3	--	1.3
Total	47	60	89	34	25.6	52.1

Source: Turkstat, Author's calculation. "Gap or Credit" refers to CPI items that are responsive to either output gap or credits or both.

6 For food and enegy items only, we estimate equations with food import prices and oil prices as controls, respectively. The details are discussed in Atuk et al. (2014).

7 The items are determined according to the significance of the coefficients of any lags of output gap or credit measures based on t-statistics calculated with heteroscedasticity corrected standard errors.

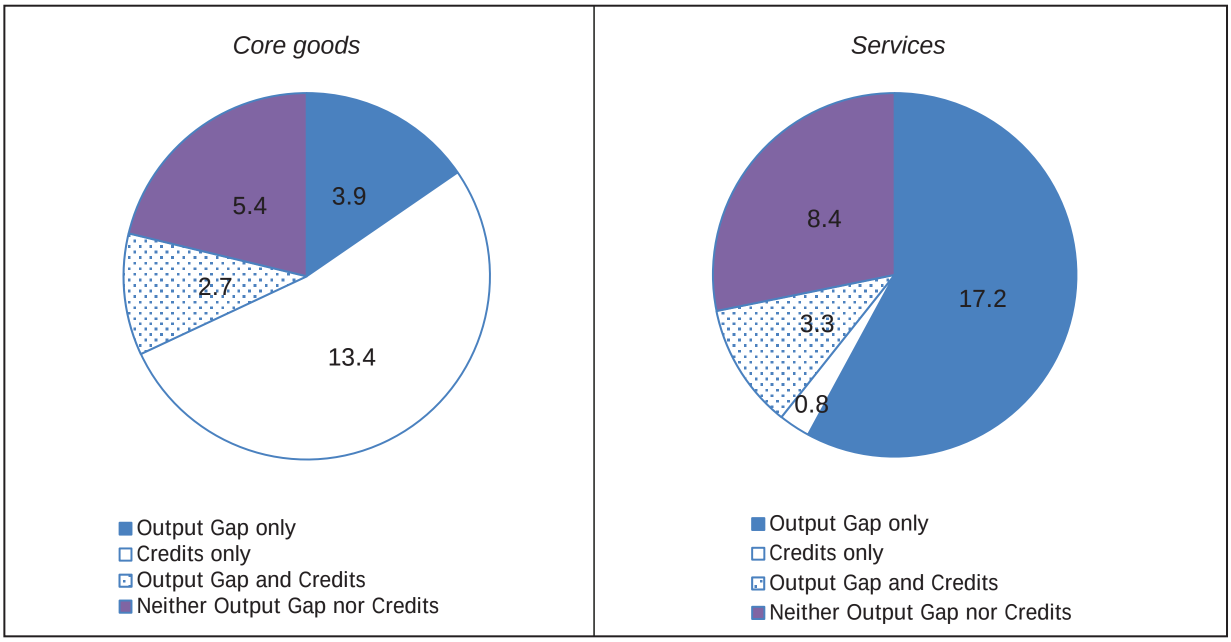
The results reveal that the sectoral decomposition of the sensitivity is quite diverse. While credit-sensitive items mostly belong to core goods, items sensitive to output gap are mostly located under services. Among the output gap-sensitive items, those in core goods group constitute 7% of the CPI, while those in services dominate 21% of the CPI. On the contrary, among the credit-sensitive items, those in core goods and services sum to 16% and 4% of the CPI respectively (Table 2).

The mostly non-tradable nature and labor-intensive structure of services, along with the observation that current income rather than credits are influential on services prices, support the strong ties between services prices and economic activity. On the other hand, considering items like durable goods and clothing whose demand can be shifted in time, the core goods prices are highly sensitive to financial conditions such as borrowing rate, maturity length and number of instalments. Therefore, credit developments play an important role for core goods inflation.

When it comes to evaluating the potential impact area of monetary policy, the question of how sensitive inflation is to intermediate targets of economic growth and credit growth through policy measures gains utmost importance. Our results show that 89 out of 152 items are sensitive to either output gap or credits, whose share in CPI is 52% (Table 2). Hence, reconciling two "demand measures", we may see that around one half of the CPI can be influenced through output gap or credits. Nonetheless, this result by no means suggests that the effect of monetary policy on inflation is limited. The index compiled by the items that are not sensitive to demand measures, co-moves with import prices in domestic currency (Appendix-Figure 1). Thus, policies aiming to stabilize exchange rate will also add to the effectiveness of monetary policy. One immediate implication of such a finding is the need for reducing the pass-through of exchange rate on prices.

Central banks follow several measures of core inflation. One of the mostly referred core indicators in Turkey is the SCA-I index which covers only services and core goods[8]. One of the motivations for following core inflation measures relates to the issue of influence. It is generally accepted that monetary policy is more effective on core measures. In this direction, our findings show that while only one half of the CPI is sensitive to output gap or credits, almost 3/4th of SCA-I can be influenced directly by the monetary policy. As seen in Figure 1, 8% of the services and 5% of the core goods, which corresponds to 1/4th of SCA-I, are not influenced by neither output gap nor credits.

Figure 1: Share of Services and Core Goods Sensitive to Output Gap and Credits in CPI (%)

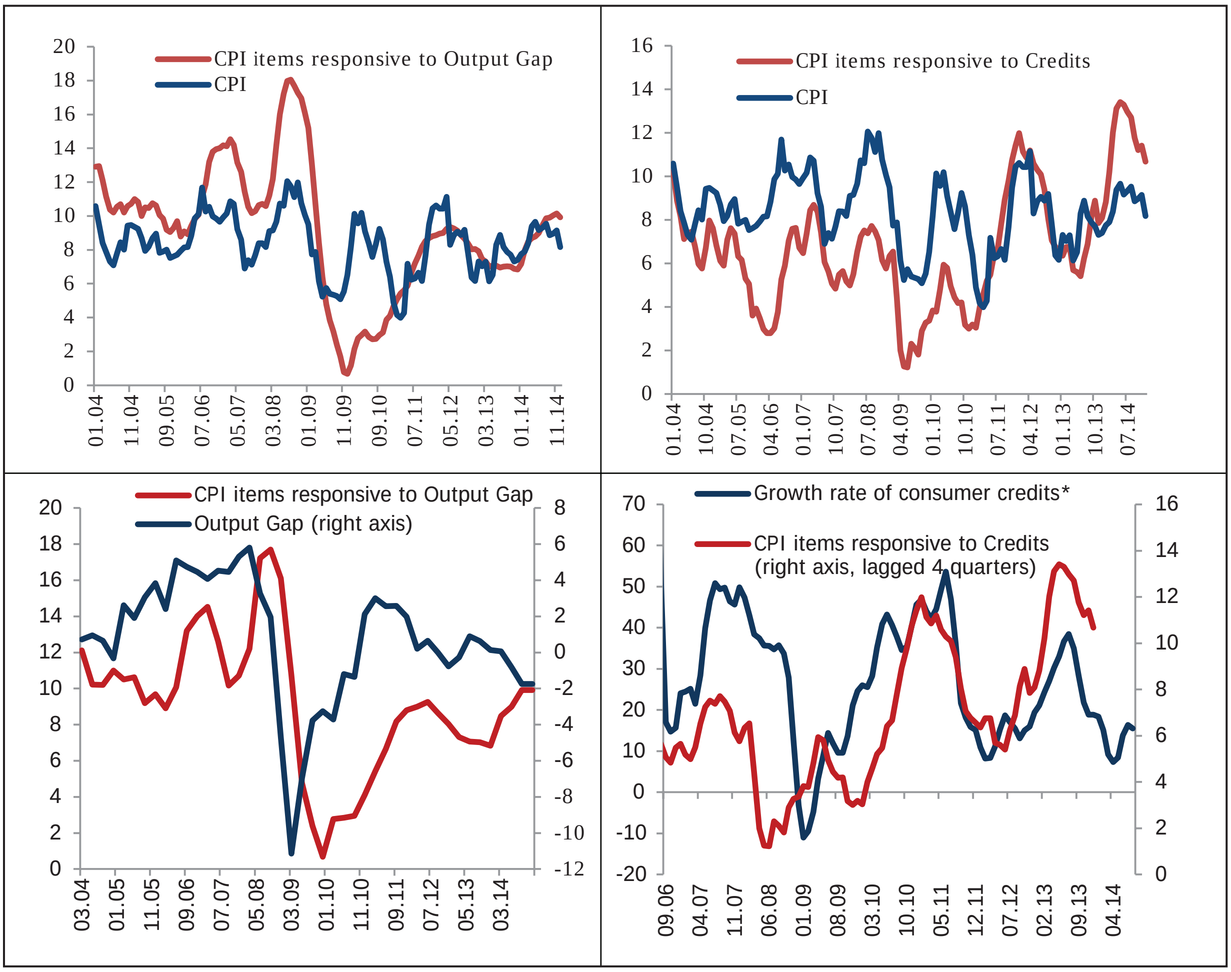


Source: Turkstat, Author's calculation.

[8] The SCA refers to Special CPI Aggregates. Those indicators are published by Turkstat along with headline CPI.

Another important result coming out of the empirical analysis directs the attention to the lag structure of the effects of the monetary policy. Due to various nominal rigidities in the economy, the transmission mechanism of monetary policy works with a certain lag. Moreover, the lag structure may differ with the choice of policy instruments. The upper panel of Figure 2 pictures the "CPI items sensitive to output gap" and "CPI items sensitive to credits" with headline inflation. Meanwhile, the lower panel depicts those two aggregate series with output gap and credit growth respectively. The difference in terms of lags is evident from the figures. While the impact of output gap on inflation is observed in 2 to 4 quarters; the effect of credit growth on inflation is observed with a much higher lag (4 to 6 quarters). This finding complies with prior evaluations of credits effecting spending and production patterns with a considerable lag.

Figure 2: Index of Items Sensitive to Output Gap and Credits (Annual % Change)



Credit growth rate is calculated as the annualized 13-week moving average rate. Source: CBRT, Turkstat, Authors' calculation.

The aggregate index of CPI items sensitive to credits is more related with domestic demand conditions. Meanwhile output gap, by definition, inherits information about foreign demand as well. Therefore, the course of both CPI items sensitive to output gap and CPI items sensitive to credits can differ dramatically on those grounds. For instance, in mid-2008, the accelerating foreign demand and hiked commodity prices fed into annual inflation of CPI items sensitive to output gap. In the meantime, slowing down domestic demand kept inflation of CPI items sensitive to credits at mild levels. On the other hand, the upward trend observed in both series in early 2007 may be attributed to exchange rate hike. Recently, the exchange rate effect is identifiable in both series despite moderate aggregate demand conditions. Those periods highlight the need for identification of foreign and domestic demand conditions along with the impact of the exchange rate.

Finally, the result also shed light on the construction of core inflation measures. For instance, most widely used exclusion based measures may fail to satisfy the criteria for core inflation measures as items under general headings like food and energy are excluded. However, the results reveal that more efficient core measures can be constructed incorporating items that are influenced by monetary policy actions through output gap and credit channels.

4. Conclusion

The question of how responsive CPI sub-indices are to medium target variables (economic activity, credits, exchange rate, etc.) that may be controlled by monetary policy is very important regarding the functioning of the transmission mechanism. Recently, CBRT devised new tools for controlling the credit growth in addition to interest rate policy. Therefore, in order to determine the stance of the monetary policy, in addition to output gap, financial indicators like credit conditions are also be followed closely. In this perspective, it is important to determine what items of the CPI actually responsive to output gap and credits. We find that one-fourth of the CPI is affected by credits, while one-third of the CPI is influenced by output gap. We also provide different aggregate price indices that would facilitate the reading of inflation developments in relation with interim monetary policy targets. Finally, the results also shed light on possible directions for the improvement of core inflation measurement.

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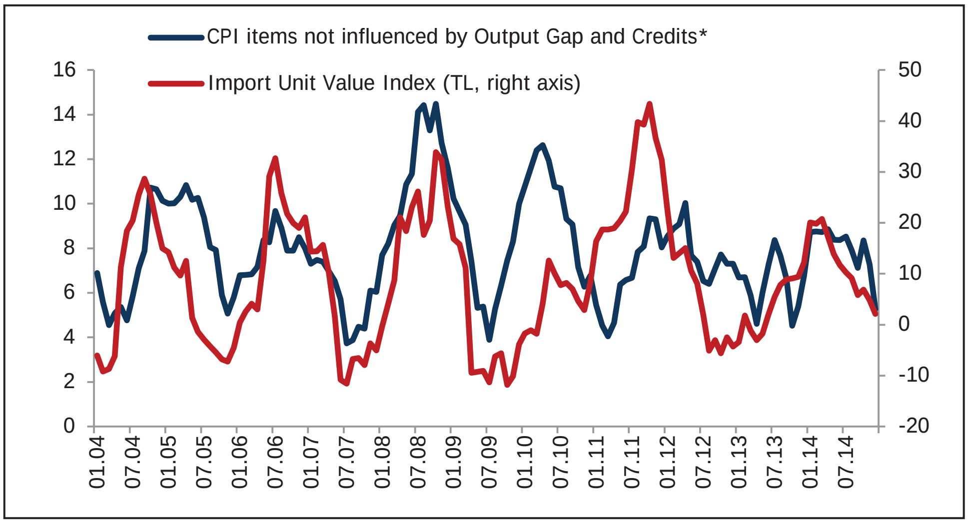
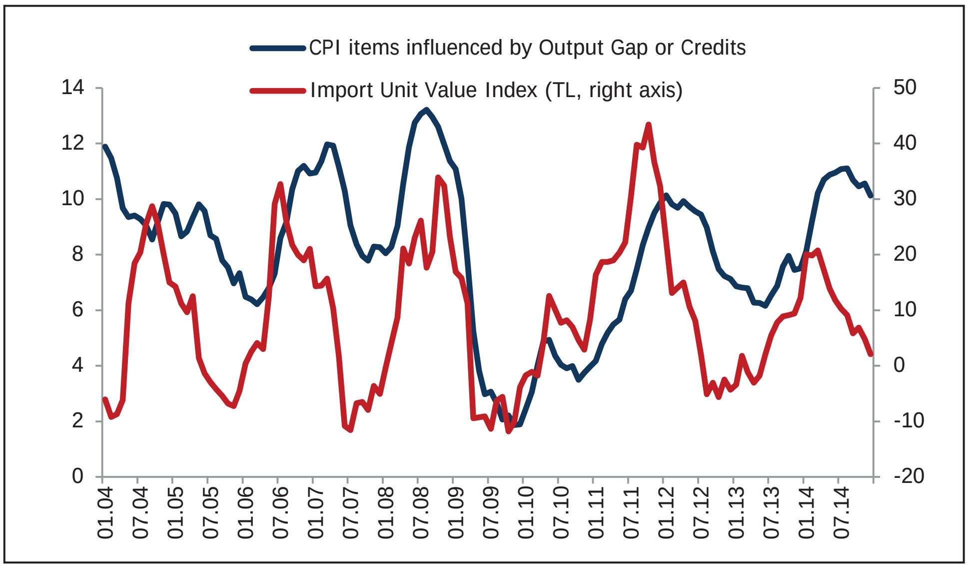
Appendix

App. Figure 1: CPI Items not Influenced by Output Gap or Credits vs. Import Prices (Annual % Change)



** The index "CPI items not influenced by output gap and credits" excludes fresh fruits and vegetables and tobacco products.*
Source: Turkstat, Authors' calculation.

App. Figure 2: CPI Items Influenced by Output Gap or Credits vs. Import Prices (Annual % Change)



Source: Turkstat, Authors' calculation.

UDC 338.31]:338.57:336.748.12}:303.22(4-672EU)

IS THERE A HARROD-BALASSA-SAMUELSON EFFECT PRESENT IN THE DATA? NEW QUARTERLY PANEL DATA EVIDENCE FROM 25 EUROPEAN COUNTRIES

Črt Lenarčič*

Abstract

The Harrod-Balassa-Samuelson phenomenon describes the relationship between productivity and price inflation within different sectors in a particular economy, where the sectoral productivity differential stands as one of the drivers for the price inflation. The Harrod-Balassa- Samuelson effect could therefore represent an additional inflation source of the economy. From the economic policy perspective it is important to address this issue, in order to contain inflation sufficiently low, especially for the likes of future EU and later on euro area accession countries, which are obliged to satisfy the Maastricht criterion. Using the fixed effects panel data model with *AR(*1) disturbances we analyse whether the Harrod-Balassa-Samuelson hypothesis is confirmed by the sectoral price inflation and the labour productivity growth data in 25 European countries. There is evidence suggesting that the faster labour productivity growth in the tradable sector relative to the non-tradable sector leads to a price inflation in the non-tradable sector considering the 25 European countries, which confirms the Harrod-Balassa-Samuelson hypothesis.

JEL Classification Numbers: C12, C23, E31

Keywords: Harrod-Balassa-Samuelson effect, productivity, inflation, fixed effects panel data model

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1 Introduction

The Harrod-Balassa-Samuelson (HBS) phenomenon describes the relationship between productivity and price inflation within different countries, regions or sectors. The productivity biased approach in explaining the purchasing power parity is an old and a well known idea. Harrod (1933), Balassa (1964) and Samuelson (1964) independently developed and formulated this idea into a model, now known as the Harrod-Balassa-Samuelson model[1]. The idea behind it is that the growth in the productivity of a tradable sector implies a rise in the wages for the tradable sector, which in turn causes also a rise in the wages for the non-tradable sector. The rise of the wages is consequently accompanied by an increase in the real exchange rate or just in the price inflation of the non-tradable sector (depending also on the exchange rate regime). Betts and Kehoe (2008) studied the relationship between the real exchange rate and the relative price of non-tradable to tradable goods. Their conclusion is that the relation between the two variables is stronger in an intense trade environment, such as the North Atlantic Free Trade Agreement (NAFTA), the free trade agreement in the European Union (EU) or the free trade agreement of the Association of Southeast Asian Nations (ASEAN). Therefore, the assumption is, that the relationship between the relative growth in the productivities of the non-tradable to tradable sector and the relative price of non-tradable to tradable goods is relatively straightforward using sectoral data for European countries.

To test the HBS hypothesis one needs to differentiate between and formulate separate entities from the data. These entities can in general represent countries, regions, or in my case sectors. For the purpose of dividing the two sectors onto a tradable sector and a non-tradable sector we follow the De Gre- gorio, Giovannini and Wolf's (1994) methodology by using the ratio of exports to total production to define both sectors. In order to do that we include and combine the NACE Revision 2 10-sector breakdown statistical classification time series data of economic activities, which provides data on labour productivity and price levels across the two sectors, and the export data given in the input-output tables, which are available at the World Input-Ouput Database (WIOD). Additionally, we exclude those sectors from the analysis, which are not distinctively tradable or non-tradable.

By obtaining the relevant tradable and non-tradable data for further analysis and by using the fixed effects panel regression methodology the presence of the HBS phenomenon is empirically tested. The estimated fixed effects panel regressions suggests that the HBS effect is confirmed by the tradable and non- tradable quarterly data of the 25 European countries in the Q1/2001-Q4/2013 period. The HBS effect is stronger considering only the transitional or accession countries in the analysis. Furthermore, HBS effect is also present using only the western European countries in the analysis, however the presence of the HBS effect is smaller. The presence of the HBS effect is also tested by employing data from the precrisis period, i.e. Q1/2001-Q2/2008 period, and data from the Q3/2008-Q4/2013 crisis period. As expected the HBS effect is larger in the transitional countries in the precrisis period compared to western countries, where the HBS effect is not statistically significant. Further on, the presence of the HBS effect is also confirmed by constructing a *vis-à-vis* model.

In section 2 a quick review of the HBS related literature is presented and discussed. In section 3 short dataset description is presented and the classification of the economic activities into a tradable and non-tradable sector is defined. Based on the latter, sectoral price indices and time series of sectoral labour productivity growths are obtained. The HBS model is described in section 4, while the estimation process and results are presented in section 5. Conclusions are presented in the section 6.

[1] or shorter Balassa-Samuelson model.

2 What has been done so far — the empirical work and issues related to Harrod-Balassa- Samuelson hypothesis

Despite treating the HBS theory as an old idea, in which the sectoral productivity differential is seen as the driver for the price inflation in the non-tradable sector (Harrod, 1933; Balassa, 1964; and Samuelson, 1964), the empirical testing of the HBS effect only became more popular in recent years with the advances in econometric methods as well as with the availability of new (or additional) time series data. The availability of new data series was largely due to the establishment of the EU and later on on its enlargement process together with advances and convergence of methodologies in collecting data by the national statistical offices. At the same time, addressing the HBS issue became relevant from the economic policy perspective in trying to identify the different sources of inflation and consequently containing the inflation low with the appropriate economic policy tools. This was (and still is) especially important for the future EU accession countries, which are obliged to satisfy the Maastricht criterion of low and stable inflation.

In their comprehensive survey working paper, Tica and Družić (2006) gathered empirical evidence considering the HBS effect. They pointed out that most of the empirical work support the HBS effect. Especially strong evidence comes from the work based on the cross-section empirical studies, similar to Balassa's (1964) work. A large number of the papers focuses on studying the magnitude of the HBS effect in accession countries in the EU. Cihak and

Holub (2001) studied the presence of the HBS effect in the Czech Republic *vis-à-vis* EU countries, while allowing for differences in structures of relative prices. Jazbec (2002) considers Slovenia as the HBS case of an accession country, while Dedu and Dumitrescu (2010) tested the HBS effect using Romanian data. Papers, as from Cipriani (2000), Coricelli and Jazbec (2001), Halpern and Wyplosz (2001), Arratibel, Rodriguez-Palenzuela and Thimann (2002), Wagner and Hlouskava (2004), Mihaljek and Klau (2008), consider a larger accession country panel. Some of the work focuses also on emerging economies. Jabeen, Malik and Haider (2011) tested the HBS hypothesis on Pakistani data, while Guo and Hall (2010) tested HBS the effect on Chinese regional data.

During the course of empirical testing of the HBS hypothesis, the models became more complicated. Rogoff (1992) was the first to implement a general equilibrium framework, with which the demand side of the economy within the HBS theory was introduced. This opened the possibility to further investigate the effects of relative productivities of production factors and the effects of the demand side of the economy on price levels[2]. However, Asea and Men- doza (1994) concluded that the proof of the HBS theory within a framework of general equilibrium cannot reliably asses the relationship between output per capita and domestic relative prices. In other words, conclusions regarding the HBS theory from cross-country analyses can only be conditionally accepted since it is difficult to account for cross-country trend deviations from purchasing power parity (PPP). Even more, Bergin, Glick and Taylor (2004) showed that the relationship between output per capita and domestic relative prices had historically oscillated too much that HBS theory could be proved by cross- section empirical studies. In order to test the HBS theory their suggestion is that it should be tested with the sector-specific analysis.

As databases, especially in Europe, became more complete, new data also made it possible to study the HBS effect between individual tradable and non- tradable sectors of a particular economy. Since it is difficult to clearly divide tradable and non-tradable commodities in the real world, some of the early papers tried to identify the tradability/non-tradability of commodities. Officer (1976) proposed that manufacturing and/or industry combine a tradable sector, while the services represent the non-tradable sector. De Gregorio, Gio- vannini and Wolf (1994) used a ratio of exports to total production of each sector to define both sectors.

[2] For instance, Mihaljek (2002) concluded that the HBS effect can have an important policy implications for the EU accession countries in order to satisfy the Maastricht inflation criterion. To further investigate Mihaljek's point, Masten (2008) constructed a two-sector dynamic stochastic general equilibrium model whether the Maastricht inflation criterion could be threatened by the HBS effect. Further on, Natalucci and Ravenna (2002) compared the magnitude of the HBS effect within different exchange rate regimes in the general equilibrium model, while Restout (2009) allowed for varying markups in its general equilibrium framework.

As was mentioned above, most studies of the HBS effect use datasets from the accession or transition European countries. The biggest setback of all empirical studies of the HBS effect is that most of the studies suffer from data measurement problems, especially from the problem of short time series availability. Researchers tried to compensate the short time series problem by pooling data from different accession/transition economies (for instance De Broeck and Slok, 2001). Others, such as Fischer (2002), Lojschova (2003), and Egert (2002), used fixed effects panel data regressions in trying to bypass the short time series problem as well as the possible data-pooling problem. For the same reason Sonora and Tica (2009) use the panel cointegration tests. We use the fixed effects panel data model by considering an autoregressive process of order 1 in the error term (AR(1)), which leads to more-efficient parameter estimates in longer panels.

Another problem that could arise is the decision regarding the choice for productivity proxy in the HBS model. In the empirical studies mostly total factor productivity (TFP) or average productivity of labour are used. Marston (1987), De Gregorio et al. (1994), De Gregorio and Wolf (1994), Chinn and Johnston, (1997), Halikias, Swagel and Allan (1999), Kakkar (2002), and Lo- jshova (2003) use total factor productivity as a productivity proxy, while due to the lack of data on TFP many others, such as Coricelli and Jazbec (2001), Zumer (2002), use average productivity of labour. In comparison between total factor productivity and average productivity of labour, the argument against the use of the average productivity of labour is, that it is not completely clear, if the average labour productivity should be regarded as a reliable indicator for representing a sustainable productivity growth, which has a long term effect on the economy (De Gregorio and Wolf, 1994). However, according to Canzoneri, Cumby and Diba (1999) the argument against the TFP is that the TFP is a result of a possibly unreliable data collection of sectoral capital stocks comparing to data collection of sectoral employment and sectoral gross value added, especially in the case of the shorter term frequencies. Sargent and Rodriguez (2000) also concluded that if the intent of the research is to examine trends in the economy over a period of less than a decade or so, labour productivity would be a better measure than total factor productivity. According to Kovacs (2002), another setback of using TFP is that, during the catch-up faze the capital accumulation intensifies faster in the transition/accession countries than in the developed countries, due to the lower starting point in fundamentals of transition/accession countries. Therefore we might overestimate the HBS effect. Listing some of the arguments against using the TFP, we consider the average labour productivity as a productivity proxy in the model. The detailed construction of the average labour productivity proxy is presented in the next chapter.

3 Data and sectoral definition

The dataset is comprised from the quarterly sectoral data available at the European Commission's statistical database Eurostat[3] for 25 European countries[4]. The time series data spans from Q1/2001 till Q4/2013 and includes sectoral labour input in the form of number of workers, and sectoral gross value added data. The dataset also includes sectoral production and services price indices data. After obtaining productivity indicator and sectoral inflation data we can proceed with the estimation of the HBS model.

3.1 Tradability of the sectors

To begin with, the tradability of the sectors has to be defined first. Since it is difficult to clearly divide tradable and non-tradable commodities in the real world, several papers try to identify the tradability/ non-tradability of commodities. Officer (1976) proposed that manufacturing and/or industry activities combine a tradable sector, while the services represent the non-tradable sector. For the purpose of the division of commodities into tradables and non- tradables, De Gregorio, Giovannini and Wolf (1994) take a step further and use a ratio of exports to total production to define both sectors. Their division threshold

[3] Available at the European Commmission's statistical database site http://epp.eurostat.ec.europa.eu/portal/page/portal/eurostat/home/

[4] Austria, Belgium, Bulgaria, Cyprus, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Latvia, Lithuania, Luxembourg, Netherlands, Portugal, Romania, Slovakia, Slovenia, Spain, Sweden, and United Kingdom.

is set to 10 percent, where the sector is defined as tradable, whether the ratio of exports exceeds the 10 percent threshold and the sector is defined as non-tradable, whether the ratio of exports does not exceed the 10 percent threshold.

Data on exports is extracted from the input-output tables available at the World Input-Output Database (WIOD). We use a standard ISIC/NACE Revision 2 "high-level' aggregation categories, which are used for reporting data from the System of National Accounts (SNA) for a wide range of countries. A 10-sector breakdown is presented in the Table 1.

Table 1: NACE Revision 2 10-sector classication of economic activities

NACE Rev. 2	10-sector breakdown description
A	Agriculture, forestry and shing
B,C,D and E	Manufacturing, mining and quarrying and other industry
F	Construction
G, H and I	Wholesale and retail trade, transportation and storage, accommodation and food services
J	Information and communication
K	Financial and insurance activities
L	Real estate activities
M and N	Professional, scientic, technical, administration and support services
O, P and Q	Public administration, defence, education, human health and social work services
R, S, T and U	Other services

Source: European Commission's statistical database Eurostat.

In order to divide the 10 sectors between tradable and non-tradable sectors, we use a similar approach as De Gregorio et al. (1994). However, we exclude those sectors from the analysis, which are not distinctively tradable or non-tradable, i.e. if their ratio of exports to total production oscillates around the 10 percent threshold too much[5]. This means that sectors such as agriculture, forestry and fishing (A), information and communication (J), financial and insurance activities (K), professional, scientific, technical, administration and support services (M and N) are excluded from the analysis. Therefore manufacturing, mining, quarrying and other industry (B, C, D and E), wholesale, retail, transportation, storage, accommodation and food services (G, H and I) are treated as tradable sectors, while construction (F), real estate activities (L), public administration, defence, education, human health, and social work services (O, P and Q), and other services (R, S, T and U) are treated as non-tradable sectors[6].

3.2 Sectoral inflation and productivity

Based on quarterly data available from the European Commission's statistical database Eurostat and the consideration of the classification of economic activities into a tradable and non-tradable sector (as defined in Table 1) supported by time-varying sectoral gross value added weights expressed in millions of euros in 2005, the aggregate price indices and percentage changes in prices for the tradable and non-tradable sectors are obtained for the 25 European countries.

The same principle applied to divide economic activities into tradable and non-tradable sectors is also used for dividing sectoral average labour productivities. Even though it is not completely clear how good a proxy labour productivity is for TFP, we use the sectoral gross value added per employee to obtain

5 A sector is treated as tradable if its ratio of exports exceeds the 10 percent threshold for at least 75 percent of time using the WIOD data for all 25 European countries and timespan from 2000 till 2011. The same principle is applied for the definition of a non-tradable sector. A sector is treated as non-tradable if its ratio of exports is under the 10 percent threshold for at least 75 percent of time using the WIOD data for all of the 25 European countries and timespan from 2000 till 2011.

6 For robustness checks we also consider more loose cases where sectors (B, C, D, E) and (G, H, I) are treated as strictly tradable and all other sectors as non-tradable, and vice versa, where sectors (F), (L), (O, P, Q) and (R, S, T, U) are treated as strictly non-tradable and all other sectors as tradable.

a relevant sectoral labour productivity indicator on a quarterly frequency basis[7]. This is done by expressing sectoral gross value added, expressed in millions of euros in 2005, per number of employees, expressed in 1000's. As was done for the tradable and non-tradable price indices the trad- able and non-tradable labour productivity indicators are also supported by time-varying sectoral gross value added weights. Based on the obtained absolute labour productivity indicator we calculate the year on year growths in the labour productivity on a quarterly frequency basis. As total factor productivity, using output per employee as a productivity proxy has its own setbacks. Due to changes in capital intensity labour productivity might deviate from total factor productivity. However, Kovacs (2002) noted that as the capital intensifies faster in the transition/accession countries in comparison to the developed countries during the catch-up faze due to the lower starting point in fundamentals of transition/accession countries, we might overestimate the HBS effect with total factor productivity methodology.

4 The theoretical price-productivity relationship

Balassa (1964) and Samuelson (1964) independently constructed a theoretical benchmark model of the real exchange rate determination meaning that faster productivity growth in the tradable than in the non-tradable sector leads to a decline in the price of tradable goods relative to the price of non-tradablegoods or affects the nominal exchange rate. Firms of both sectors are subject to the following Cobb-Douglas production functions (1928)

$$y_{T,t} = A_{T,t} l_{T,t}^{\alpha_T} k_{T,t}^{1-\alpha_T}, \tag{1}$$

and

$$y_{N,t} = A_{N,t} l_{N,t}^{\alpha_N} k_{N,t}^{1-\alpha_N}, \tag{2}$$

where y denotes output, l labour input, k capital input, and A is the productivity indicator. Subscripts T and N denote tradable and non-tradable goods, whereas a denotes output elasticity of capital and 1 — α denotes output elasticity of labour input. Under the assumption of perfect competition in capital and labour markets the wages in the two sectors will be equal to the marginal product of labour

$$w_{T,t} = p_{T,t} \alpha_T A_{T,t} \left(\frac{k_{T,t}}{l_{T,t}} \right)^{1-\alpha_T}, \tag{3}$$

and

$$w_{N,t} = p_{N,t} \alpha_N A_{N,t} \left(\frac{k_{N,t}}{l_{N,t}} \right)^{1-\alpha_N}, \tag{4}$$

If we take into consideration the case of a small open economy with perfect labour mobility, nominal wages in the tradable and non-tradable sectors will be the same, $w_{T,t} — w_{N,t}$. Combining and rearranging (3) and (4) we get

$$\frac{p_{N,t}}{p_{T,t}} = \frac{\alpha_T A_{T,t} \left(\frac{k_{T,t}}{l_{T,t}} \right)^{1-\alpha_T}}{\alpha_N A_{N,t} \left(\frac{k_{N,t}}{l_{N,t}} \right)^{1-\alpha_N}} = \frac{\alpha_T \frac{y_{T,t}}{l_{T,t}}}{\alpha_N \frac{y_{N,t}}{l_{N,t}}}. \tag{5}$$

7 Since the total factor productivity (TFP), which is defined as the portion of output not explained by the amount of inputs used in production and measures the efficiency and the intensity of input utilization in the production process (Comin, 2008), has its setbacks, especially in the form of the availability of relevant data. Input-output tables are at best available on yearly frequency and the time series of the input-output data for European countries are relatively short.

Log-differentiating (5) leads to

$$\widehat{p}_{N,t} - \widehat{p}_{T,t} = \widehat{\widehat{a}}_{T,t} - \widehat{\widehat{a}}_{N,t}, \tag{6}$$

where $\widehat{\widehat{a}}_{T,t} = \log \frac{y_{T,t}}{l_{T,t}}$ and $\widehat{\widehat{a}}_{T,t} = \log \frac{y_{T,t}}{l_{T,t}}$. The intuition behind the equation (6) is that there is a positive link between faster productivity growth in the tradable sector relative to the non-tradable sector and the growth of non- tradable prices relative to prices of tradable goods. This is known as the Harrod-Balassa-Samuelson effect.

5 Estimation and results

This section provides the estimated model and the regression results. So far, the HBS hypothesis was tested in numerous papers using a range of econometric methods. Several papers, such as Bahmani-Oskooee (1992), Bahmani- Oskooee and Rhee (1996), Chinn (1997), Halikias et al. (1999), Deloach (2001), Taylor and Sarno (2001), Egert (2002), consider different cointegra- tion methodologies, i.e. E/G method (Engle and Granger, 1987) or Johansen and Juselius' (1990) method. Generalized method of moments was used by Halpern and Wyplosz (1998), and Arratibel et al. (2002), Hsieh (1982) uses instrumental variable method, while De Broeck and Slok (2001), and Fischer (2002) tested the HBS effect with the autoregressive distributed lag method. Despite the wide variety of different econometric methods applied in testing the HBS hypothesis, the most widely used techniques are still the OLS and GLS estimation methods (Canzoneri et al., 1999; Coricelli and Jazbec, 2001; Halpern and Wyplosz, 2001; Egert et al., 2003). With the availability of additional sectoral data the fixed effects panel data model was introduced (Fischer, 2002; Lojschova, 2003).

Bearing in mind the constructed dataset of the price inflation and labour productivity growth of the tradable and non-tradable sectors from the 25 European countries, the reason of choosing the fixed effects panel data model is straightforward. It allows us to control for variables that cannot be observed and is suitable for multilevel modelling. In our case the unobservables are the country-specific differences. In other words it accounts for country-individual heterogeneity.

5.1 Estimated models 5.1.1 The baseline model

This estimated model follows a de Gregorio et al. (1994) type of a model setting with an internal identification of the HBS effect, however it is upgraded into a 25 European country fixed effects panel data model allowing for an AR(1) process in the error term. The model can be written as

$$p^{NT}_{i,t} = c_i + \beta_1 A^{TN}_{i,t} + \beta_2 gdp_{i,t} + \beta_3 exp_{i,t} + \beta_4 gov_{i,t} + \beta_5 cap_{i,t} + \nu_t + u_{i,t}, \tag{7}$$

where variable p^{NT}_{it} is the relative price of non-tradable sector goods to tradable goods, $p_{N,i,t} — P_{T,i,t}$, β*'s* measure the impact of the independent variables, while i is the number of countries entering the estimation process. The variable ci therefore captures the country-specific effects, vt is the vector of the year dummies and $u_{i,t}$ is the error term following the *AR(1)* process $u_{i,t} = pu_{i,t-1} + \varepsilon_{i,t}$.

In comparison to *i.i.d.* error terms in other panel data models, the assumptions of the *AR(1)* process in the disturbances will potentially lead to more efficient parameter estimates, especially in longer-panel models. The key explanatory variable, A^{TN}_{it}, includes differences in productivity growths between the tradable and non-tradable sector, i.e. the HBS effect, $\alpha_{T,i,t} — \alpha_{N,i,t}$.

We also include other driver variables with which unwanted deviations in the estimation model are explained. Changes in the real GDP, $gdp_{i,t}$, capture deviations related to changes in economic activity, while the variable $exp_{i,t}$ tries to capture deviations related to changes in exporting behaviour of countries, which can have a direct effect on the price inflation in the short run, giving more pressure on the price inflation of the tradable goods. Furthermore, also changes in the government spending, $gov_{i,t}$, are considered. Changes in government spending will cause prices to rise, especially the rise in the inflation of the non-tradable goods, since most of the government spending falls on service activities. Variable $cap_{i,t}$ captures changes in gross fixed capital formation and steers away the effects of unwanted deviations in the capital formation from the labour productivity growth.

The usage of other explanatory variables is quite common in the HBS literature. In general, we could separate these variable into several groups. First, the income type of proxies include the GDP and consumption related variables (De Gregorio et al., 1994; Halpern and Wyplosz, 2001; Frensch, 2006; Gubler and Sax, 2011) in order to trap the regular economic fluctuations in a model. Second, proxies related to government activities such as the government spending (De Gregorio et al., 1994; Fischer, 2002; Sonora and Tica, 2009; Gubler and Sax, 2011), wages (Halikias et al., 1999; Mihaljek and Klau, 2008), deficit (Arratibel et al., 2002) and debt (Rogoff, 1992) mostly influence the deviations of the non-tradable side of the economy. On the other hand, the third type of variables, proxies related to the openness of the economy capture the unwanted deviations of the tradable sector. These proxies are usually changes in exports of goods and services (Halpern and Wyplosz, 2001), current account (Gubler and Sax, 2011), export-to-GDP ratios, and other economy openness measures. In the last group all the other explanatory variables are included, such as exchange regime dummies (Arratibel et al., 2002), changes in capital accumulation, and oil-price changes (Chinn, 1997; Chinn and Johnston, 1997; De Broeck and Slok, 2001).

5.1.2 The *vis-à-vis* model

With the baseline model we try to show the statistical presence of the HBS effect in the selected panel of the 25 European countries. By using most of the available 25 European countries data and dividing them into transitional/accession economies[8] and euro area economies[9] (EA12) we build a visa-vis model with EA12 as a numeraire country. The *vis-à-vis* model is written as

$$p_{i,t}^{NT} - p_t^{EA12} = c_i + \beta_1(A_{i,t}^{TN} - A_t^{EA12}) + \beta_2(gdp_{i,t} - gdp_t^{EA12}) + \beta_3(ex_{i,t} - exp_t^{EA12}) + \beta_4(gov_{i,t} - gov_t^{EA12}) + \beta_5(cap_{i,t} - cap_t^{EA12}) + \beta_6 fx_{i,t} + \nu_t + u_{i,t}, \tag{8}$$

where $u_{i,t}$ is the error term following the $AR(1)$ process $u_{i,t} = \rho u_{i,t-1} + \varepsilon_{i,t}$. The variable c_i captures the transitional-country-specific effects with i number of transitional economies, while the variable v_t is the vector of the year dummies. Dependent variable $p_{i,t}^{NT} - p_t^{EA\,12}$ is the relative price of non-tradable sector goods to tradable goods vis-à-vis transitional countries versus EA12 countries. The same principle applies for the independent variables. The variable $A_{i,t}^{TN} - A_t^{EA\,12}$ is the difference between the transitional and EA12 countries in the relative productivity growths of the tradable and non-tradable sector. The other driver variables, changes in the real GDP, $gdp_{i,t} - gdp_t^{EA12}$, government spending, $gov_{i,t} - gov_t^{EA12}$, exports, $ex_{i,t} - exp_t^{EA12}$, and gross fixed capital formation, $cap_{i,t} - cap_t^{EA12}$, are also treated in the relativistic fashion, the transitional countries vis-à-vis the EA12.

In the vis-à-vis model we also relax the Betts and Kehoe (2008) assumption regarding the strong relationship between the real exchange rate and the relative price of non-tradable to tradable goods in an intense trade environment. By doing this, a new independent variable is considered in the model, the nominal exchange rate, $fx_{i,t}$.

5.2 Panel data regression results 5.2.1 The baseline model results

The baseline model results are discussed in this subsection. All regressions include an intercept and year dummy variables, while fixed effects for each of the 25 European countries are considered. To minimise or partly exclude the seasonality problem and the possible one-off effects in the economies, which can be present in quarter on quarter growth data, the year on year data on a quarterly frequency

8 Bulgaria, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Romania, Slovenia, and Slovakia.

9 Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, and Spain.

is considered. Year on year price data enter the estimated model as the difference between the non-tradable sector and the tradable sector, while year on year labour productivity growth data enter the baseline model as the difference between the labour productivity growth in the tradable sector and the labour productivity growth in the non-tradable sector. Year on year changes in real GDP, government expenditure, exports, wages, and gross fixed capital formation are also considered in the model. Further details on the variables considered in the baseline model are presented in the descriptive statistics Table 2.

Table 2: Descriptive statistics of the baseline model variables (in %)

Variable	Obs.	Mean	Std. Dev.	Min	Max
$p_{i;t}^{NT}$	1300	1.426	4.613	-29.368	24.009
$A_{i;t}^{TN}$	1300	1.313	12.354	-69.615	57.521
$gdp_{i;t}$	1300	1.989	4.063	-18.993	13.644
$exp_{i;t}$	1300	4.816	8.858	-31.142	66.512
$gov_{i;t}$	1300	1.635	4.492	-23.806	45.074
$cap_{i;t}$	1300	1.673	11.638	-43.380	51.595

The regression results of the baseline model give an interesting conclusion. In all of the 5 regressions, when assuming that the change in the difference of productivity growths is instantaneously reflected in the change of the price ratio between the non-tradable and tradable sector, the coefficient of the difference of productivity growths reflecting the HBS effect, $a_{T,i,t} - a_{N,i,t}$, is positive and statistically significant (see the results in Table 3).

In order to exclude any unwanted deviations in the model other explanatory variables are also considered. In the regression 1, the changes in real GDP, changes in exports and changes in government spending are included in the analysis. In the regression 2, changes in gross capital formation are additionally considered, however it does not change the magnitude of the HBS effect. In both regressions the changes in GDP and exports are, as the HBS effect, statistically significant and are in line with the theory discussed above. The rise in real GDP increases the relative price of non-tradable sector goods to tradable goods, while the rise in exports puts more pressure on the price inflation of the tradable goods, which in turn decreases the relative price of non-tradable sector goods to tradable goods. Government spending and gross capital formation seem not to have a particular effect looking at the results in regressions 1 and 2.

While transition countries were experiencing higher growth due to the catching-up of the developed economies, as expected, the HBS effect is stronger when including only data from the 9 transition economies[10]. A similar result was shown by Garcia-Solanes and Torrejon-Flores (2009) the emerging economies tend to have a stronger HBS effect than the developed ones. The HBS effect is therefore much smaller when including only the data from the 16 western economies[11], as shown in the regression 4, and including only the data 12 euro area countries - EA12[12], as shown in the regression 5. Looking at the other driver variables one can observe that changes in the government spending in western economies and consequently in EA12 countries also matters in explaining the relative price movement. The rise in the real GDP and government spending increase the relative price of non-tradable sector goods to tradable goods, while the rise in exports puts more pressure on the price inflation of the tradable goods, which in turn decreases the relative price of non-tradable sector goods to tradable goods.

[10] Bulgaria, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Romania, Slovakia, and Slovenia.

[11] Austria, Belgium, Cyprus, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden United, and Kingdom.

[12] Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, and Spain.

Table 3: Results of the baseline model

Regressions no.	1	2	3	4	5
$A_{i;t}^{TN} = a_{T;i;t} - a_{N;i;t}$	.0450***	.0449***	.0548***	.0259**	.0283**
the HBS effect	(.0105)	(.0105)	(.0186)	(.0125)	(.0136)
$gdp_{i;t}$	.2899***	.2871***	.3098***	.2214***	.2980***
	(.0548)	(.0602)	(.0962)	(.0691)	(.0868)
$exp_{i;t}$	-.0948***	-.0947***	-.1151***	-.0628***	-.0644**
	(.0180)	(.0180)	(.0304)	(.0217)	(.0292)
$gov_{i;t}$	.0231	.0231	.0073	.0684**	.1134**
	(.0206)	(.0206)	(.0339)	(.0276)	(.0463)
$cap_{i;t}$		.0016			
		(.0149)			
Constant, year dummies	Yes	Yes	Yes	Yes	Yes
Sector effects	Fixed	Fixed	Fixed	Fixed	Fixed
Number of countries	25	25	9	16	12
Observations	1275	1275	459	816	612
R^2	.1827	.1826	.2172	.1641	.1808
$corr(u_i; Xb)$	.0099	.0101	-.0235	-.0330	-.0825
$\rho_{AR(1)}$)	.5880	.5879	.4986	.6810	.6786

Note: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

In Table 4, the same panel regression estimation exercise is carried out using only the precrisis data, i.e. data before Q2/2008. The results give similar conclusions as results presented in the Table 3, however the HBS effect seems to be smaller. Again, dividing the panel into transition (regression 8), western (regression 9) and EA12 countries (regression 10), it is clear that the HBS effect is even stronger looking at the transitional countries panel (despite being smaller on aggregate level), mainly due to the catching-up phase before the global financial crisis. However, during this period the HBS effect was almost non-existing in the western and EA12 economies. Also the effects of the other explanatory variables are in line with the theory discussed above, however the effects of the GDP and government spending changes are not statistically significant. The main driver amongst the other explanatory variables are therefore changes in exports, especially in the transition countries.

Table 4: Results of the baseline model - precrisis data

Regressions no.	6	7	8	9	10
$A_{i;t}^{TN} = a_{T;i;t} - a_{N;i;t}$	.0296**	.0296**	.0609**	-.0123	-.0211
the HBS effect	(.0141)	(.0141)	(.0291)	(.0133)	(.0141)
$gdp_{i;t}$	.1554*	.1555*	.2354	.1126	.1709*
	(.0855)	(.0901)	(.1597)	(.0849)	(.1012)
$exp_{i;t}$	-.0831***	-.0832***	-.1136***	-.0288	.0134
	(.0217)	(.0218)	(.0373)	(.0246)	(.0331)
$gov_{i;t}$	-.0083	-.0082	-.0081	.0235	.0580
	(.0264)	(.0265)	(.0436)	(.0371)	(.0668)
$cap_{i;t}$		-.0000			
		(.0210)			
Constant, year dummies	Yes	Yes	Yes	Yes	Yes
Sector effects	Fixed	Fixed	Fixed	Fixed	Fixed
Number of countries	25	25	9	16	12
Observations	725	725	261	464	348
R^2	.0773	.0685	.1449	.0235	.0339
$corr(u_i; Xb)$	.0687	.0685	-.0573	-.0151	-.0999
$\rho_{AR(1)}$)	.5170	.5156	.4251	.6358	.6535

Note: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

On the other hand, in Table 5, the panel regression estimation exercise is carried out using only data from the crisis period, i.e. data from Q3/2008 onward. The HBS effect seems to be particularly stronger looking at the panels of the western (regression 14) and EA12 countries (regression 15), despite the presence of the relatively strong HBS effect in transition countries (regression 13). During the crisis period also the significance of other driver variables grows, meaning that the relative price of the non-tradable and tradable sector are much more affected by the changes in real GDP, exports and government spending.

Table 5: Results of the baseline model - crisis period data

Regressions no.	11	12	13	14	15
$A^{TN}_{i;t} = a_{T;i;t} - a_{N;i;t}$ the HBS effect	.0599***	.0602***	.0469*	.0913***	.1214***
	(.0160)	(.0160)	(.0240)	(.0243)	(.0269)
$gdp_{i;t}$	.2956***	.3068***	.2956*	.1922	.3060*
	(.0909)	(.0984)	(.1509)	(.1276)	(.1608)
$exp_{i;t}$	-.1339***	-.1341***	-.1527***	-.1029**	-.1462***
	(.0337)	(.0338)	(.0585)	(.0405)	(.0524)
$gov_{i;t}$	.0882**	.0883**	.0992	.0993**	.1610**
	(.0385)	(.0386)	(.0723)	(.0449)	(.0716)
$cap_{i;t}$		-.0066			
		(.0223)			
Constant, year dummies	Yes	Yes	Yes	Yes	Yes
Sector effects	Fixed	Fixed	Fixed	Fixed	Fixed
Number of countries	25	25	9	16	12
Observations	525	525	189	336	252
R^2	.1291	.1284	.1212	.1656	.2191
$corr(u_i ; Xb)$	.0538	.0524	.0132	.1613	-.0319
$\rho_{AR(1)}$)	.6253	.6254	.5570	.6882	.6750

Note: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

5.2.2 The *vis-à-vis* model results

The *vis-à-vis* model results are discussed in this subsection. As in the baseline model, all of the regressions include an intercept and year dummy variables, while fixed effects for each of the 9 European transition/accession countries are considered. Similarly, the seasonality problem is minimised by using the year on year data on a quarterly frequency. Year on year relative non-tradable to tradable prices data enter the *vis-à-vis* model as the difference between the transition countries and the numeraire country, i.e. EA12. As for the prices, the relative tradable to non-tradable productivities data enter the model as the difference between the transition countries and the EA12. Differences between the transition countries and the EA12 are also considered for year on year changes in real GDP, government expenditure, exports, wages, and gross fixed capital formation. On the other hand, the nominal exchange rate enters the model only as the year on year change. Further details on the variables considered in the baseline model are presented in the descriptive statistics Table 6.

Table 6: Descriptive statistics of the *vis-à-vis* model variables (in %)

Variable	Obs.	Mean	Std. Dev.	Min	Max
$p^{NT}_{i;t} - p^{EA12}_{t}$	468	1.093	5.489	-14.777	17.359
$A^{TN}_{i;t} - A^{EA12}_{t}$	468	3.010	13.706	-59.519	56.924
$gdp_{i;t}$	468	2.413	3.928	-14.850	11.707
$exp_{i;t}$	468	4.329	7.675	-26.269	55.634
$gov_{i;t}$	468	-.071	5.748	-25.752	42.257
$cap_{i;t}$	468	4.642	12.631	-39.055	46.746
$fx_{i;t}$	468	.654	5.722	-13.989	33.511

The estimation results of the *vis-à-vis* model are presented in Table 7. In regressions 16 and 17 data from the whole period Q1/2001-Q4/2013 is analysed, while in regressions 18 and 19 the precrisis data is used, and in regressions 20 and 21 the crisis period data is used. Taking into account the *vis-à-vis* relationship between the transitional and EA12 countries, one can observe that the HBS effect is quite constant irrespective of the period, precrisis or crisis. Looking at the other explanatory variables, the inflation differential was mostly (negatively) driven by changes in exports in the precrisis period. During the crisis period, i.e. after the Q3/2008, changes in the real GDP and the (nominal) appreciation of the local transition country currency against the euro affect the relative prices.

Table 7: Results of the vis-a-vis model

Regressions no.	16	17	18	19	20	21
$A^{TN}_{i;t}$ - A^{EA12}_{t}	.0583***	.0565***	.0560*	.0554*	.0575**	.0541**
the HBS effect	(.0188)	(.0189)	(.0291)	(.0291)	(.0241)	(.0242)
$gdp_{i;t}$ gdp^{EA12}_{t}	.3301***	.2789**	.1890	.1673	.4325***	.3456**
	(.1032)	(.1101)	(.1654)	(.1701)	(.1522)	(.1648)
$exp_{i;t}$ exp^{EA12}_{t}	-.1054***	-.1010***	-.1227***	-.1198***	-.1087	-.1018
	(.0331)	(.0333)	(.0388)	(.0392)	(.0709)	(.0710)
$gov_{i;t}$ gov^{EA12}_{t}	.0025	.0023	-.0040	-.0040	.0834	.0811
	(.0343)	(.0341)	(.0441)	(.0442)	(.0730)	(.0723)
$fx_{i;t}$	-.0999	-.1033	.0566	.0563	-.3581***	-.3556***
	(.0694)	(.0699)	(.0866)	(.0869)	(.1183)	(.1182)
$cap_{i;t}$ cap^{EA12}_{t}		.0350		.0222		.0504
		(.0264)		(.0398)		(.0356)
Constant, year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Sector effects	Fixed	Fixed	Fixed	Fixed	Fixed	Fixed
Number of countries	9	9	9	9	9	9
Observations	459	459	262	262	189	189
R^2	.1534	.1458	.1405	.1329	.1509	.1440
$corr(u_i; Xb)$	-.0849	-.0838	-.0777	-.0986	-.1511	-.1474
$\rho_{AR(1)}$)	.4968	.5060	.4208	.4228	.5391	5516

Note: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

5.2.3 Policy implications

Even though we have shown the statistical significance and presence of the HBS effect, the results suggest that it does not play a significant role in determining the inflation differential, in the baseline case and — pfA12 in the *vis-à-vis* case. However, these results can have important economic policy implications, especially for the likes of future EU and later on euro area accession countries, which are obliged to satisfy the Maastricht criterion of low and stable inflation. In order to satisfy this criterion, these countries would have to implement adequate measures in their efforts of containing inflation and not attribute the overall inflation to the HBS effect, which is relatively smaller than previously thought. Similar conclusions regarding relatively small impact of the HBS effect were made by Subasat (2010) and Podkaminer (2003). However, addressing the HBS issue could still be beneficial for other transition or emerging economies in choosing the appropriate economic policy tools.

6 Conclusions

With the advances in econometric methods as well as with the availability of new (or additional) time series data the empirical testing of the Harrod- Balassa-Samuelson effect became more popular in recent years. Using the fixed effects panel regression estimation we show that the Harrod-Balassa-Samuelson hypothesis is confirmed by price and labour productivity data at a quarterly frequency considering the 25 European countries. The Harrod-Balassa- Samuelson effect is particularly stronger including only the data from transition/accession economies in comparison to the Harrod-Balassa-Samuelson effect which includes the data from developed economies. Using only the precrisis data the Harrod-Balassa-Samuelson effect is even stronger in transition/accession countries, while the Harrod-Balassa-Samuelson effect is almost non-existing in the developed countries. The Harrod-Balassa-Samuelson effect is also confirmed in *vis-à-vis* type of model setting, where the Harrod-Balassa- Samuelson effect is tested for the transitional countries and euro area countries as a numeraire country.

Despite the statistical significance and presence of the Harrod-Balassa- Samuelson effect, the estimation results suggest that it does not play a major role in determining the inflation differential. These results can have important economic policy implications, as for EU and euro area accession countries and also other transition or emerging economies. Based on these results, these countries would have to deploy other economic policy measures or tools to contain the overall inflation, since the Harrod-Balassa-Samuelson effect is relatively smaller than the previous wide-spread perception.

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FORENSIC ANALYSIS OF CREDIT ACTIVITY IN CROATIA*

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Abstract

The main purpose of this paper is to identify the determinants and evolution of credit demand and supply of households and corporates in Croatia for the period from 2000 until the second quarter of 2014. Approach is based on the switching regression framework and results for the most recent crisis period are cross checked with the information obtained from the bank lending survey. The results show there are factors that limit the possibility of intensifying credit activity for the households and corporates both on the supply, and demand side, with more pronounced drag coming from the subdued demand.

(please do not quote)

JEL classification: E44, G21, G28

Key words: credit supply, credit demand, households, corporates, Croatia

* Views presented in this paper are the views of the authors and do not necessarily express views of the Croatian National Bank.

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1 Introduction

In the period after the escalation of the global financial crisis, credit activity in Croatia slowed down considerably compared to the pre-crisis period. Household loans recorded negative annual rates of change since mid-2009, while corporate loans started to decrease at the end of 2012. Although the literature does not provide unequivocal proof of the nature of the relationship between credit activity and economic growth or pace of recovery, the identification of determinants of credit supply and demand is important for understanding the capacity and scope of monetary policy to influence loan dynamics. The fact that credit activity slowed down despite the expansive monetary policy which resulted in high banking system liquidity, raises the question as to whether the reasons for such developments lie on the supply side because. For instance, if banks are less inclined to offer loans due to more tightened lending standards. Or the issue is on the demand side as a result of negative current real developments, structural problems in the balance sheets of the private sector and pessimistic expectations regarding future economic developments.

The main purpose of this paper is to identify the determinants and evolution of credit demand and supply of households and corporates in the period from the first quarter of 2000 until the second quarter of 2014, using the switching regression framework. This paper broadens the paper by Čeh et al. (2011) on the disequilibrium in the market of total domestic and foreign loans, with the analysis for individual corporate and household sectors. Additionally, this analysis has also been expanded by findings on credit supply and demand obtained from the bank lending survey (survey), which the Croatian National Bank (CNB) has conducted since 2012. The results show there are factors that limit the possibility of intensifying credit activity for the households and corporates both on the supply, and demand side, with more pronounced drag coming from the subdued demand.

The paper is divided in five parts. After the introduction part, a brief overview of credit activity in Croatia from 2000 to 2014 is given, and in the third part the methods used for the analysis of the determinants of credit supply and demand are presented – the credit market disequilibrium model and the bank lending survey. The fourth part presents the results of the estimated model and describes the development of surplus or deficit credit supply through time, while for the recent period these results have been supplemented by the findings from the survey. The paper ends with concluding considerations that have been put in the context of the scope of monetary policy to influence the revival of credit activity in the current phase of the economic cycle.

2 Credit activity in Croatia

The period from 2000 to 2014 in Croatia was marked by the credit cycle that in its first part was characterised by excessive credit growth, followed by its deceleration. A similar pattern marked many other Central and Eastern European (CEE) as a result of the process of convergence of these economies towards the old EU Member States before the crisis, or the "sobering up" after the escalation of the global financial crisis. During this period, the determinants of credit supply and demand overlapped and changed courses of action and CNB measures also played an important role in loan dynamics.

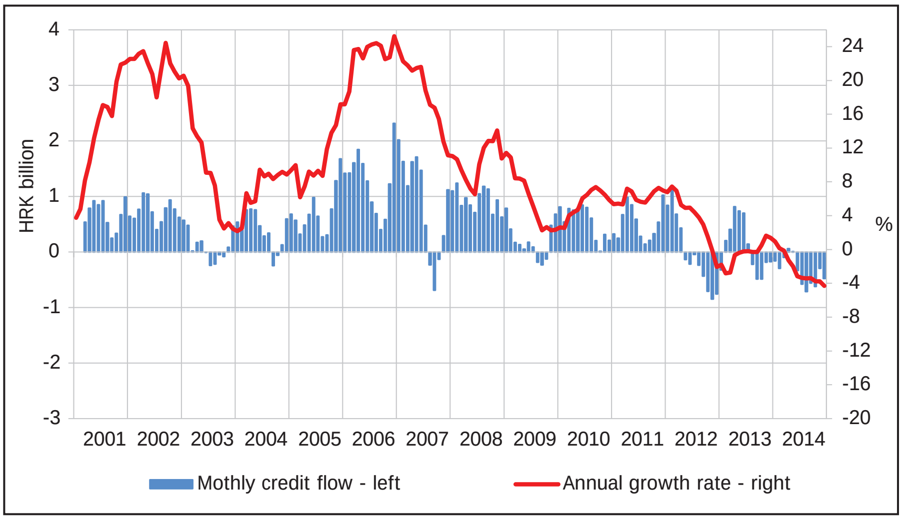
Credit growth in Croatia in the pre-crisis period was supported by strong foreign capital inflows that financed the increase in domestic consumption and resulted in the accumulation of macroeconomic imbalances because of large deficits in the balance of payments current account and a strong increase in external debt. A significant part of these inflows referred to funds received from parent banks of the largest domestic banks, which represented cheap sources of financing for their daughter banks in Croatia. In such conditions, credit supply was abundant, and lending conditions relaxed, which was greatly contributed by a strong competition among banks to conquer market shares. Domestic credit demand was strong, and, in addition to increase in consumption, it was also stimulated by the boom in the real estate market. Consumer confidence was high, and expectations of growth in future income were anchored. At the same time, the CNB conducted an active countercyclical policy directed at the decrease in profitability of foreign sources of financing and discouraging expansive credit supply. Ljubaj (2012) confirms that measures which prior to the crisis limited excessive credit growth, really influenced loan dynamics, especially for the households.

However, after Croatia slumped in the prolonged recession from 2009 onwards, the developments in the credit market changed considerably. The share of non-performing loans started to increase fast, the profitability of banks started to decline, and the private sector, burdened by high indebtedness and balance sheet weaknesses became a less desirable debtor so that the aversion of banks towards assuming new risks increased significantly. By contrast, the decline in economic and especially in investment activity dampened corporate credit demand, with a more expressed effect of unfavourable trends in the labour market on households, which started a continuous process of deleveraging. With the beginning of the crisis, the nature of monetary policy changed, so that the CNB began to encourage credit programmes and relaxed monetary policy instruments with the aim of increasing monetary system liquidity and improve domestic financing conditions, thus enabling the recovery of lending.

Looking at the credit developments for the corporate sector, the growth in corporate loans was most prominent from 2001 to 2002, and in the period from 2006 to 2007, while the rates of change in loans were positive until the end of 2012, despite the absence of economic recovery (Figure 1). The average annual growth rate of corporate placements from 2001 until the end of 2008 was almost 20%, and after the escalation of the crisis it recorded a multiple decrease and stood at about 2%. During the entire observed period, corporates also significantly relied on foreign financing, most often from banks related to domestic banks in terms of ownership, in particular when CNB measures for limiting domestic credit growth were in force. Although the number of corporates that only have domestic debt considerably exceeds the number of external indebted corporates (over 30 thousand compared to about a thousand[1]), the balance of external debt is higher than all domestic corporate loans, which confirms a strong influence of foreign financing on the corporate credit market in Croatia.

However, despite the above-average increase in corporate loans during the 2010 and 2011, in comparison with the majority of the CEE countries, the deleveraging of this sector has intensified recently. Highest drop in corporate loans has been recorded in 2014 (- 4%) suggesting that deleveraging is maybe in the early stage of progress. The continuation of this process might be another factor weighing on the fragile recovery of the Croatian economy, which additionally motivates the analysis of the factors having impact on the credit supply and demand for this sector.

Figure 1 Placements to corporate sector in Croatia

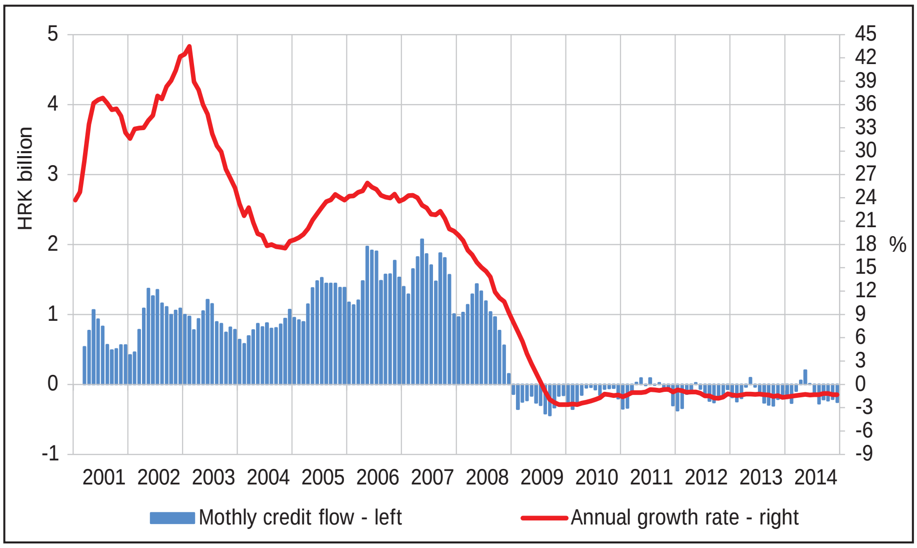


Note: Data are adjusted for exchange rate movements and one-off effects, including loan sales, bank bankruptcies, methodological changes and the government assumption of the shipyards' debt. Data for monthly credit flows are 3-month moving averages.
Source: CNB

[1] CNB's Bulletin No. 205, Box 3 An overview and structure of the debt of non-financial corporations in 2013

The divergence of credit dynamics before the crisis and during the crisis is more expressed in households than in corporates (Figure 2). From 2001 until the end of 2008, household placements grew on average over 25%. Lending grew strongly in this period in conditions of real growth in disposable income of households, and a considerable easing of price and non-price related lending conditions. Simultaneously, a relatively faster growth in total debt from income considerably worsened the indicators of their indebtedness and increased the risks related to the possible increase in the debt repayment burden.

Figure 2 Placements to household sector in Croatia



Note: Data are adjusted for exchange rate movements and one-off effects, including loan sales, bank bankruptcies and methodological changes. Data for monthly credit flows are 3-month moving averages.
Source: CNB

The decline in real incomes during recession spurred the adjustment of household balance sheets. In 2009, credit growth was halted, and since mid-2009, households have been constantly deleveraging at a very stable rate, so that household placements decrease by about 1.5% annually. In the environment of the delayed recovery of the Croatian economy, unfavourable developments in the labour market and negative expectations concerning future developments and income, it is uncertain when this trend might be stopped or reversed. Also, the CNB's analysis of the process of deleveraging of households indicates at the need for a further short-term reduction in the debt burden, where the model suggests that the intensification of economic activity and increased consumer confidence would open the possibility for this sector to take on further debt (CNB, FSR 13).

3 Methods of identification of determinants of credit supply and demand

This paper builds on the working paper by Čeh et al. (2011) who, using the credit market disequilibrium model on the example of Croatia, established whether the limiting factor of credit activity at a certain moment was on the side of supply or on the side of demand. This analysis was carried out for total domestic bank loans and for foreign loans to domestic sectors for the period from 2000 to 2010. In this analysis, the determinants of supply and demand are estimated specifically for the household sector and for the non-financial corporations from 2000 until the second half of 2014, and the results of the model are supplemented by the results obtained by the bank lending survey: In that way, information and findings about the development of supply and demand from the creditors' perspective are also taken into account, and at the same time, the robustness of model-evaluations is checked.

3.1 Disequilibrium model in the market of corporate and household loans

One of the ways to identify the determinants of credit supply and demand is through econometric or model approach. The credit market disequilibrium model may serve for this purpose, with the aim to determine the periods of surplus or deficit of credit supply or demand at a specific time, and to identify the factors that determine credit supply or demand by corporates and households. The model builds on the paper by Ghosh and Ghosh (1999)[2], who used the switching regression framework in the analysis of credit supply and demand developments, for which the foundation had been laid by Maddala and Nelson (1974), who proposed the maximum likelihood method for the evaluation of such models, and Quandt and Ramsey (1978) who, using the switching regression framework, viewed the periods of surplus or deficit of credit supply as two regimes with a certain likelihood of occurrence.[3]

The household and corporate loan market disequilibrium models were rated using the maximum likelihood method.[4] Using a system of simultaneous equations, the main determinants of real credit supply and demand of corporates and households were established and the periods of surplus supply or demand for each of these sectors were identified. Similar approach is applied in Everaert et al. (2015), who use such model to assess whether credit demand or credit supply was the factor constraining the evolution of actual credit in five CEE countries.

The dependent variable both in the function of credit supply and credit demand are bank loans to the corporate and household sectors. Surplus supply or demand were calculated as the difference between the estimated credit demand and credit supply, with the loans actually utilised equalling at each given moment the lower of the values between supply and demand. Data used in this analysis are quarterly data for the period from the first quarter of 2000 to the second quarter of 2014.

Accordingly, the observed credit C_t must equal the lesser of the two quantities[5], i.e.:

$$C_t = \min(C_t^d, C_t^s)\,,$$

where:

$$C_t^d = \mathbf{X}_{1t}'\boldsymbol{\beta}_1 + \varepsilon_{1t}$$

$$C_t^s = \mathbf{X}_{2t}'\boldsymbol{\beta}_2 + \varepsilon_{2t}$$

X_{1t}' represents the determinants of credit demand, X_{2t}' determinants of credit supply, $\boldsymbol{\beta}_1$, $\boldsymbol{\beta}_2$ are the parameters to be estimated, and $\varepsilon_{1t}, \varepsilon_{2t}$ are random errors.

Assuming that the errors are independent and normally distributed with variances σ_1^2 and σ_2^2, the following functions are given:

$$f_1(C_t) = \frac{1}{\sqrt{2\pi\sigma_1^2}} \exp[-\frac{1}{2\sigma_1^2}(C_t - \mathbf{X}_{1t}'\boldsymbol{\beta}_1)^2]$$

$$f_2(C_t) = \frac{1}{\sqrt{2\pi\sigma_2^2}} \exp[-\frac{1}{2\sigma_2^2}(C_t - \mathbf{X}_{2t}'\boldsymbol{\beta}_2)^2]$$

$$F_1(C_t) = \frac{1}{\sqrt{2\pi\sigma_1^2}} \int_{C_t}^{\infty} \exp[-\frac{1}{2\sigma_1^2}(C_t - \mathbf{X}_{1t}'\boldsymbol{\beta}_1)^2]dC_t^d$$

2 The same authors estimated a similar model within the *Credit Crunch or Weak Demand for Credit?* on the examples of Latvia, Hungary and Poland, published in the World Bank Report *EU 10 – Regular Economic Report* in October 2009.

3 For a detailed overview of the literature by models for the evaluation of credit market disequilibrium please see Čeh et al., 2011.

4 For a detailed overview of the model please see Čeh et al., 2011.

5 This condition helps to avoid the usual identification problems in credit market equilibrium models, given that, in each period, the volume of credit is determined by either supply or demand.

$$F_2(C_t) = \frac{1}{\sqrt{2\pi\sigma_2^2}} \int_{C_t}^{\infty} \exp[-\frac{1}{2\sigma_2^2}(C_t - \mathbf{X}'_{2t}\boldsymbol{\beta}_2)^2] dC_t^s$$

$$f(C_t^d, C_t^s) = f_1(C_t^d) f_2(C_t^s) = \frac{1}{2\pi\sigma_1\sigma_2} \exp\left(\frac{-\left(C_t^d - \mathbf{X}'_{1t}\boldsymbol{\beta}_1\right)^2}{2\sigma_1^2}\right) \exp\left(\frac{-\left(C_t^s - \mathbf{X}'_{2t}\boldsymbol{\beta}_2\right)^2}{2\sigma_2^2}\right)$$

The likelihood that an observation in a period t is determined by demand $C_t = C_t^d < C_t^s$ is given by a conditional density C_t

$$f(C_t | C_t = C_t^d < C_t^s) = \frac{\int_{C_t}^{\infty} f(C_t, C_t^s) dC_t^s}{\Pr ob(C_t^d < C_t^s)} = \frac{f_1(C_t) \int_{C_t}^{\infty} f_2(C_t^s) dC_t^s}{\Pr ob(C_t^d < C_t^s)} = \frac{f_1(C_t) \cdot F_2(C_t)}{\Pr ob(C_t^d < C_t^s)}$$

where $f(C_t, C_t^s)$ is the joint density of C_t and C_t^s. Conversely, if in a period t there is a credit crunch $C_t = C_t^s < C_t^d$, then the conditional density C_t is given by

$$f(C_t | C_t = C_t^s < C_t^d) = \frac{\int_{C_t}^{\infty} f(C_t, C_t^d) dC_t^d}{1 - \Pr ob(C_t^d < C_t^s)} = \frac{f_2(C_t) \int_{C_t}^{\infty} f_1(C_t^d) dC_t^d}{1 - \Pr ob(C_t^d < C_t^s)} = \frac{f_2(C_t) \cdot F_1(C_t)}{1 - \Pr ob(C_t^d < C_t^s)}$$

Given that in a period t the observed quantity of credit to each sector is determined by either supply or demand, the unconditional density C_t is given by

$$f(C_t | \mathbf{X}_{1t}, \mathbf{X}_{2t}) = \Pr ob(C_t^d < C_t^s) f(C_t | C_t = C_t^d) + (1 - \Pr ob(C_t^d < C_t^s)) f(C_t | C_t = C_t^s)$$
$$= f_1(C_t) \cdot F_2(C_t) + f_2(C_t) \cdot F_1(C_t)$$

on the basis of which the *LogLikelihood* of the entire sample is

$$L(\boldsymbol{\beta}_1, \boldsymbol{\beta}_2 | \mathbf{X}_{1t}, \mathbf{X}_{2t}) = \sum_{t=1}^{n} \log[f_1(C_t) \cdot F_2(C_t) + f_2(C_t) \cdot F_1(C_t)]$$

The parameter estimations for $\boldsymbol{\beta}_1$, $\boldsymbol{\beta}_2$ will represent the values maximising the *LogLikelihood.*

The *LogLikelihood* maximisation algorithm has been initiated from the starting point for $\boldsymbol{\beta}_1$, $\boldsymbol{\beta}_2$ estimated by OLS or 2SLS. Given that the series are non-stationary, the estimated results are reasonable only if the credit supply and demand determinants are cointegrated, which can be[6] proved by a test of cointegration between the model-estimated credit demand and the observed credit series and, separately, between the estimated credit supply and observed credit.

6 Please, note that the estimated supply and demand series represent a linear combination of their determinants.

3.2 Bank lending survey

Supply and demand is often analysed using data obtained by regular creditor or bank lending surveys. Usually such surveys are conducted by central banks and their results provide an important source of information on price and non-price related lending terms and conditions, as well as factors that impact credit demand. Aggregate responses provided by banks are frequently used, as well as responses at micro-level (depending on the type of the model that is estimated). In numerous papers credit supply and demand is often analysed using data from bank lending surveys. For example, Bondt et al. (2010) confirmed that the bank lending survey offers useful information to forecast loan growth, GDP and investment in the euro area, and Hempell et al. (2010) analysed the impact of supply constraints on credit developments in the euro area on the basis of the survey.

The CNB's survey is methodologically aligned with the survey conducted for the euro area by the European Central Bank, so that it covers the questions that refer to the previous quarter and to the expectations in the subsequent three months. Questions are grouped with regard to two types of banks' credit portfolio, households and corporates, and responses to them are provided by bank managers responsible for credit operations with these sectors. Lending standards include internal rules and written and unwritten criteria and reflect an individual bank's credit policy (for instance, requirements which potential clients must meet for a certain type of loan the bank is inclined to grant, collaterals which the bank is inclined to accept, etc.). Lending terms are subject to the agreement between the lender and the borrower, such as the interest rate amount and the scope of the security instrument (e.g., requests for collateral, margins on average risk loans, margins on riskier loans, fees, maturity, etc.). So far, ten surveys have been conducted in which banks, whose assets amount up to 99% of total assets of the banking system, have participated regularly.

Results are commented on the basis of the net percentage of the banks' responses weighted by the size of the banks. For lending standards, net percentage is the difference between the share of the banks that have tightened lending standards and the share of those that have eased them. A positive net percentage indicates that the share of the banks that have tightened their lending terms is higher than the share of the banks that have eased them, so that it is the case of net tightening. In the contrary case, it is net easing. Therefore, when commenting on credit demand, a positive net percentage indicates that the share of the banks responding that the demand has increased is higher than the share of the banks reporting a decline in demand, so that it is the case of net growth in demand, and vice versa.

As the CNB started conducting bank lending surveys in October 2012, the time series available is not sufficiently long to provide data capable of being used in the credit market disequilibrium model. Also, the relatively short series of banks' responses does not allow us to put the present results into a longer context (for example how much the present responses deviate from long-term averages). However, for the recent period, survey results may be combined with those from the model estimated on a much longer time sample, by which the robustness of the model results is partially checked.

4 Results of the analysis

In this chapter, the results of the estimated model of credit supply and demand determinants for the household and corporate sectors are presented, which are supplemented by recent findings of the bank lending surveys conducted so far.

4.1 Corporates

Independent variables in the estimated function of corporate credit demand, which have proved to be significant, include the nominal interest rate on corporate loans, real GDP, GDP gap, EMBI yield spread, corporate profitability and business confidence index, while the function of credit supply is best determined by the banks' credit potential[7], the difference between lending and deposit interest rates, real GDP, the credit risk indicator of banks and the development of risk premium for the country and for parent banks of the largest domestic banks. The results of the model are presented in Table 1.

The estimated model shows that higher economic activity results in the strengthening of corporate credit demand, while faster than potential GDP growth acts in the opposite direction, which can probably indicate at the increased possibility for corporate internal financing in the conditions of stronger expansion, and vice versa (Table 1). Higher corporate profitability, coupled with increased business confidence, is associated with heightened investment activity, while an increase in lending interest rate reduces credit demand. An increase in the EMBI yield spread for Croatia boosts demand since it makes the possibility of substitution of domestic loans by foreign borrowing more difficult. Not surprisingly, credit supply is positively influenced by higher economic activity, increased loan potential and greater profitability of deposit and lending operations. The results of the model show that a larger amount of partly and fully irrecoverable placements increases credit supply, which can be explained by increased efforts of banks to dilute the share of bad placements in total placements by inflows of new, recoverable loans, although it is a less likely behaviour of banks in the period of crisis. At the same time, the increase in provisioning of bad loans reduces earnings, and it can also jeopardise the capital of the banks, which may eventually limit credit supply. Higher country risk premium reduces credit supply, in the same way as higher risk premium of parent banks raises the price of capital and affects its allocation within the group.

Table 1 Results of the disequilibrium model on the market of corporate loans

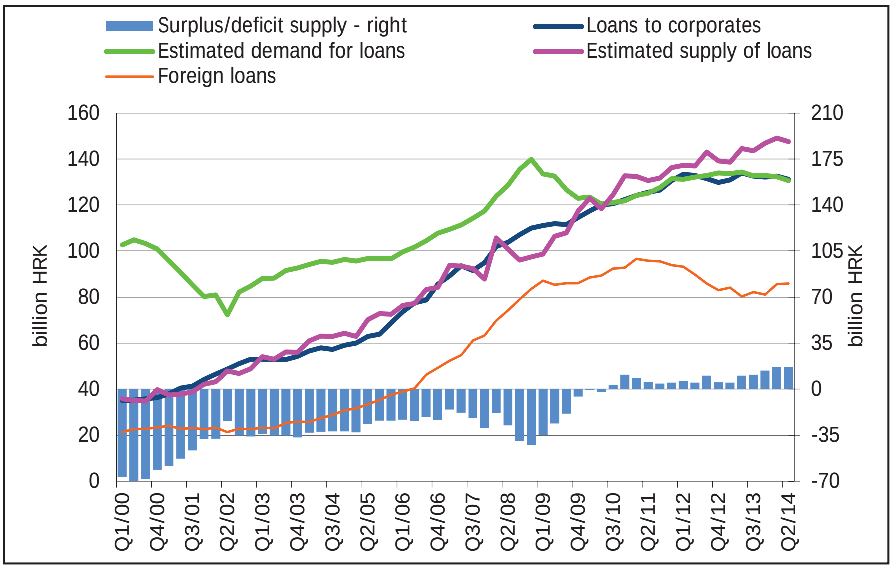
Demand	
Independent variable	
Constant	3,58**
Lending interest rate	-0,03*
GDP	1,29***
GDP gap	-0,41*
Profitability of corporate assets	0,76***
EMBI spread	0,04*
Business confidence	0,35***
Standard deviation	0,10
Supply	
Independent variable	
Constant	-3,63
Lending and deposit interest rate spread	0,01*
GDP	2,35***
Credit potential	0,21**
Non-performing corporate loans	0,47***
Loan loss provisions on corporate loans	-0,19***
EMBI spread for RC	-0,01**
Parent bank CDS	-0,10***
Standard deviation	0,04

*Note: * significant at 1%, ** significant at 5%,*** significant at 10%. Loans granted, GDP and credit potential were deflated by the consumer price index and observed in logs. GDP was seasonally adjusted and GDP gap is the difference between the original GDP series and its trend obtained by means of the Hodrick-Prescott filter. The business confidence index covering the entire observed period was obtained by constructing a new series comprised of the business confidence index of Privredni vjesnik and business expectation index.*
Source: Authors' calculation

The model estimated in this way allows for observing the evolution of credit supply and demand through time (Figure 3). The pre-crisis period was characterised by surplus demand for corporate loans relative to the supply of such loans, which can be related not only to a relatively poor availability of such loans of this sector and fast economic growth, but also to central bank measures aimed at slowing down credit activity with the aim of containing external imbalances and overheating of the domestic economy. Credit supply was primarily determined by a high level of capital inflows into the banking sector, strong economic activity and increased risk appetite as well as the probable underestimation of credit risk, as reflected in the fall in the share of non-performing household and corporate loans and lower provisioning of bad loans. A significant portion of surplus credit demand in the domestic market during that period was met by direct borrowing of corporates abroad.

[7] The credit potential is approximated by foreign liabilities and savings and time deposits.

Figure 3 Estimated supply and demand for corporate loans

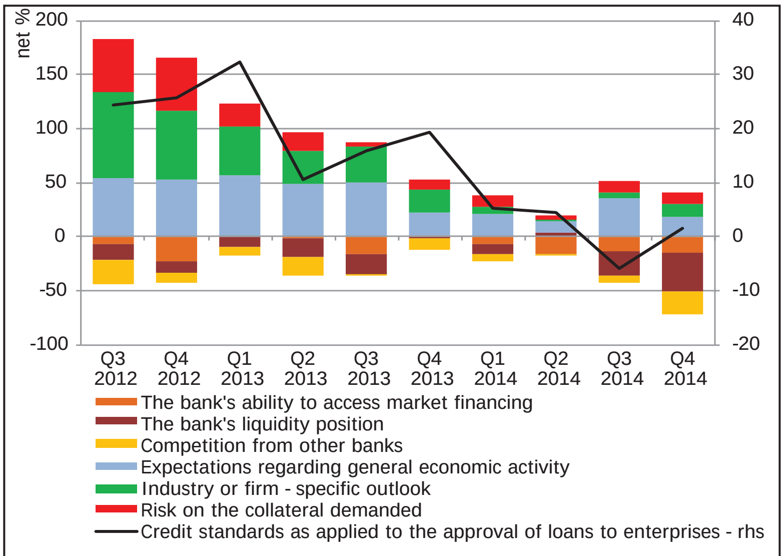


Source: Authors' calculation

The escalation of the global financial crisis in the third quarter of 2008 was followed by a considerable increase in risk premiums and slowdown in capital inflows, which affected the credit potential of banks and their inclination to grant loans and led to a temporary decrease in corporate credit supply. The freezing of the international financial markets led to a simultaneous increase in demand of corporates for domestic loans despite a mild increase in interest rates. However, credit supply stabilised quickly and continued to grow, partially supported by monetary policy relaxation which ensured financial stability in the country and the inflow of parent banks' capital into the domestic banks.

From mid-2010 until the end of the observed period credit demand mostly stagnated because of the low level of economic activity, and to a certain extent also because of the stabilisation in the international financial markets and easier access to foreign capital. For the above reasons, the period from mid-2010 onwards was marked by surplus supply of loans over demand. This, among other, resulted in a continuous process of deleveraging of domestic banks abroad and a drop in their credit potential.

Figure 4 Factors affecting credit standards as applied to the approval of loans to enterprises

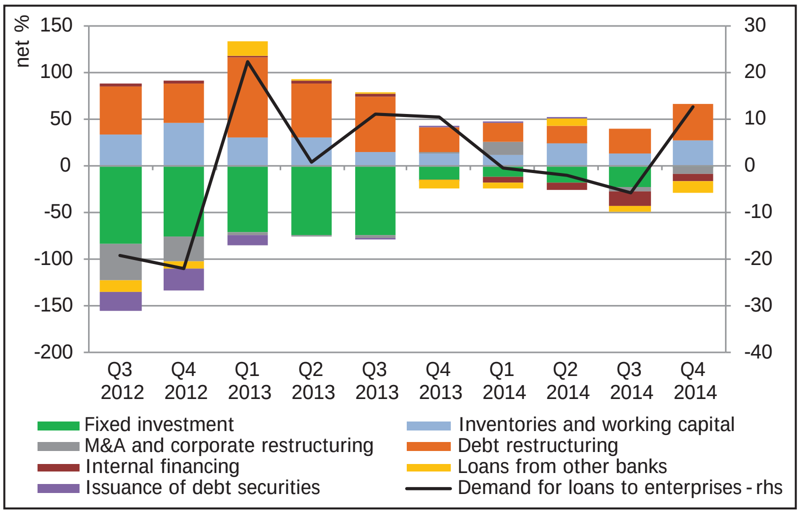


Note: The positive value shows that the factor contributes to standard tightening and the negative that it contributes to standard easing.
Source: CNB.

However, although the estimated model indicates at a surplus of supply of loans to corporates, survey results show that lending standards for corporate loans mostly tightened from the third quarter of 2012, only to ease slightly at the end of 2014 (Figure 4). In addition to the trend of easing the tightening of standards, the effect of individual factors that contributed to the subdued supply also weakened. This primarily refers to the negative expectations of general economic expectations, the outlook for industry or specific corporates and collateral risk. This is in line with the model estimation of the effect of determinants of corporate credit supply. The growth in economic activity increases credit supply, while the higher provisioning reduces it, so that it is not surprising that banks report the tightening of lending standards because of negative expectations regarding economic recovery and negative outlook for individual corporates. Collateral risk is also emphasised, which, in the end, may increase provisioning for the non-performing corporate loans. Of the factors that influence in the opposite direction, the survey shows that the liquidity of banks, the competition from other banks and the possibility of bank financing in the market make a positive contribution to the easing of standards (which can then also increase credit supply). These factors have actually been under the dominant influence of the effect of expansive monetary policy in the recent period, from which it results that conditions for bank financing in the market are not a limiting factor for credit supply.

Survey results (Figure 4) show that corporate credit demand decreased at the end of 2012, and then increased in 2013, after which less favourable developments were recorded again in addition to the recovery of demand only for the last observed period (the fourth quarter of 2014). In view of the constraints caused by the long-term recession and high debt levels and poor capitalisation of corporates, corporate demand for domestic loans was subdued, especially for new loans because the factor that mostly contributed to growth in demand was the need for debt restructuring and the financing of working capital. On the other hand, the lack of investments was the main factor of decrease in demand. Therefore, it is not surprising that in the period from 2012 to 2014 model-estimated credit demand actually stagnated, so that the gap of the surplus of supply widened additionally. It can be concluded that the factors of demand from the survey and the model estimation of demand reflect and confirm that the delayed recovery of the Croatian economy also limits the recovery of corporate credit demand.

Figure 5 Factors affecting demand for loans to enterprises



Note: The positive value shows that the factor contributes to higher demand and the negative that it contributes to lower demand.
Source: CNB.

4.2 Households

As regards the model for supply and demand in the household credit market, statistically significant on the demand side were the interest rate on household loans, personal consumption, consumer confidence index and the real net wage bill, while the credit supply function is best determined by the credit potential of banks, the difference between lending and deposit rates, real GDP, credit risk indicators of banks, real estate prices and risk premiums for parent banks of the domestic banks. As the mechanism of the effect of some variables on credit demand or supply is similar to that applicable to corporates, the effect of variables characteristic for households is described below.

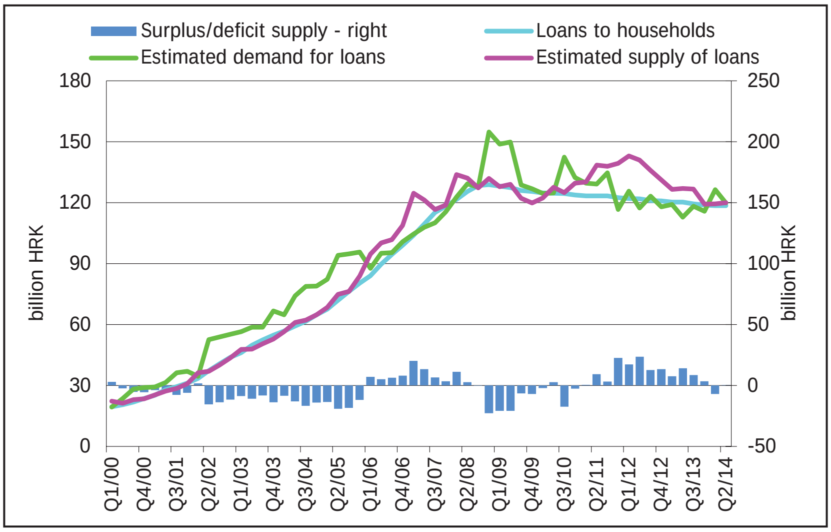
The growth in the wage bill which depends not only on wages but also on the number of employees is, not surprisingly, pushing credit demand. While an increase in real estate prices on the one hand decreases the demand for home loans, it may on the other hand also increase household borrowing capacity and the inclination of banks to grant loans to debtors due to higher collateral value. However, in both cases this coefficient was negative and did not prove significant.

Table 2 Results of the disequilibrium model on the market of household loans

Demand	
Independent variable	
Constant	-7,92**
Lending interest rate	-0,07***
GDP	1,47***
Consumer confidence	0,01***
Gross wage bill	1,32***
HREPI	-0,03
Standard deviation	0,06
Supply	
Independent variable	
Constant	-4,00***
Lending and deposit interest rate spread	0,02***
GDP	2,97***
Credit potential	0,34***
Non-performing household loans	0,31**
Loan loss provisions on household loans	- 0,21**
HREPI	- 0,05
Mother banks CDS	-0,5**
Standard deviation	0,04

*Note: * significant at 1%, ** significant at 5%, *** significant at 10%. Loans granted, GDP, wage bill and credit potential were deflated by the consumer price index and observed in logs. The GDP and the wage bill were seasonally adjusted.*
Source: Authors' calculation

Figure 6 Estimated supply and demand for household loans

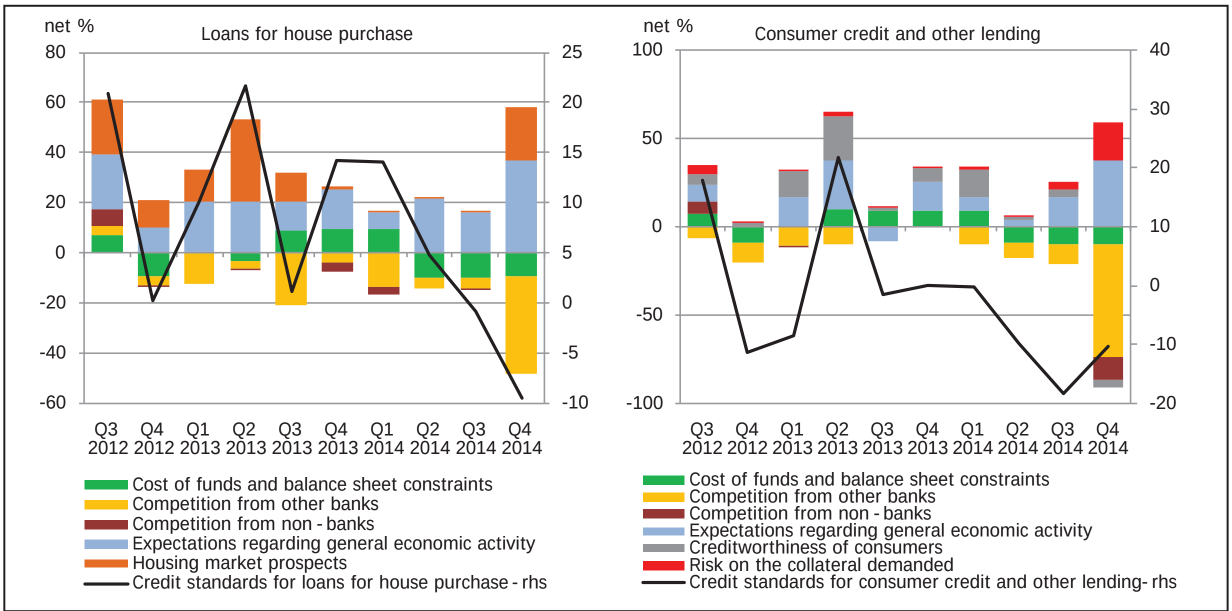


Source: Authors' calculation

Looking at the evolution of credit supply and demand for the household sector, a major part of the pre-crisis period was marked by surplus household credit demand in relation to the supply (Figure 5). This demand was mainly driven by heightened economic activity, low starting level of debt of this sector, positive developments in the labour market and increased consumer confidence, and gradual reduction in interest rates on household loans from the relatively high levels that marked the early 2000s. The reason for lower credit supply in relation to demand should also be sought in the CNB measures restricting credit activities of banks, with households, unlike corporates, having had no access to foreign sources of financing. The period of surplus supply lasted from early 2006 to mid-2007 when the CNB, in an effort to slow down growth in credit placements based on external debt growth, penalised growth in placements exceeding 12%. The beginning of the crisis in the third quarter of 2008 was followed by a brief period of increased household credit demand that exceeded the supply, which had begun to fade, while since early 2011 credit supply has generally exceeded the demand, implying that demand factors were the main drivers of household deleveraging.

The results of the bank lending survey point to tightening of lending standards for housing loans granted to households from 2012 until the third quarter of 2014, when the beginning of easing of these standards was observed. On the other hand, much more favourable developments were recorded for consumer and other loans where banks reported almost a continuous easing of standards. Negative expectations with regard to general economic trends are the main factor of the tightening of lending standards for both groups of household loans, which is confirmed by the direction of activity estimated in the household credit market disequilibrium model. Also, in the direction of tightening, the negative perspective of the real estate market for home loans and the credit capacity of the clients for consumer loans (again, it can be linked to the negative effect of value adjustments on credit supply indicated by the model) are also emphasised. The competition of other banks is the main factor that contributes to the easing of household loans and the costs of the source of financing and balance sheet restrictions have been acting in the same direction in the recent period, which is information that supplements the findings from the model estimation.

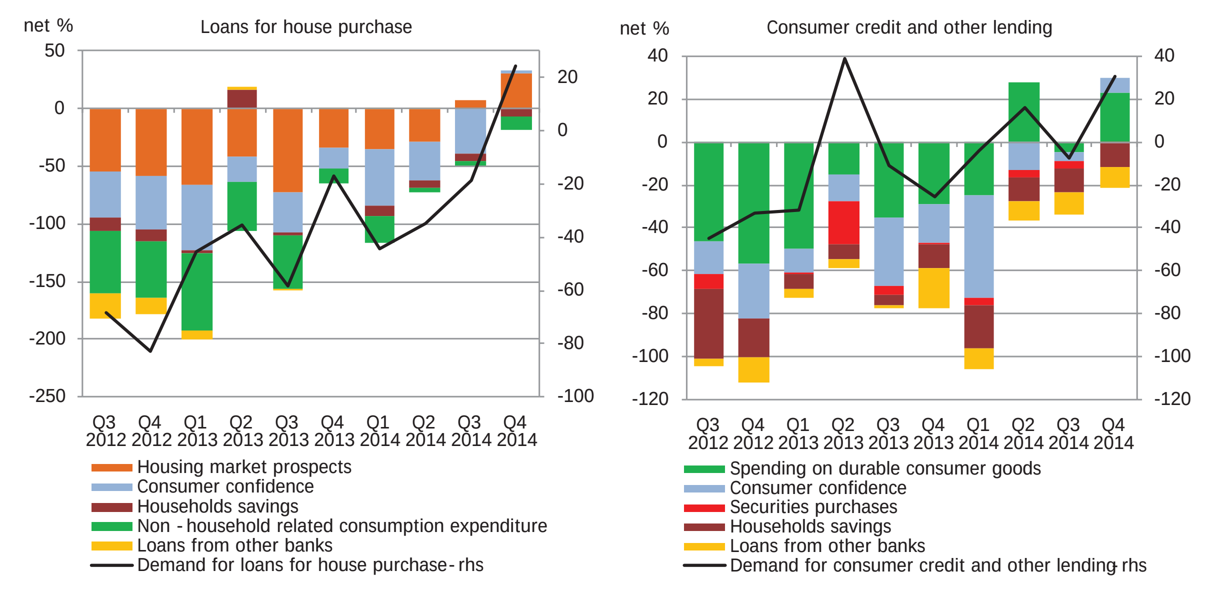
Figure 6. Factors affecting credit standards as applied to the approval of loans to households



Note: The positive value shows that the factor contributes to standard tightening and the negative that it contributes to standard easing.
Source: CNB.

The survey shows that household demand from 2012 to 2014 mostly decreased, in particular for housing loans (except in the fourth quarter of 2014), but also for consumer loans with sporadic oscillations. In general, household demand developments were less favourable than corporate demand developments, and these were most unfavourably affected by consumer confidence, household consumption, the perspectives of the real estate market and housing savings. All of this together can easily be linked to the variables and directions of the effects from the model estimation (especially GDP and consumer confidence), which again confirms that the results of the survey, which present the creditor's view on demand and supply, supplement the findings from the model estimation. Here, too, the conclusion can be drawn that without economic recovery it will be difficult to realise the recovery of credit demand, which, in the case of households, is especially subdued.

Figure 7. Factors affecting demand for loans to households



Note: The positive value shows that the factor contributes to higher demand and the negative that it contributes to lower demand.
Source: CNB.

5 Conclusion

The understanding of determinants and evolution of credit supply and demand is crucial for the analysis of the scope of monetary policy measures to influence credit activity. Although interrelations of various factors are puzzling, it seems that postponed recovery and weak growth prospect are probably most dominant factors influencing sluggish credit developments in Croatia.

The estimated model and jointly estimated credit supply and demand functions, expanded by findings from the bank lending survey which increase the robustness of the model evaluation, has shown that the main determinants of corporate and household credit demand are greatly linked to the domestic macroeconomic environment. In the absence of the recovery of economic activity and consumer and business confidence it is difficult to expect significant improvements in the domestic credit markets. The reversal of negative economic developments might also have a positive effect on the inclination of credit institutions to offer loans, which, in addition to the already low interest rates, might result in favourable non-price related financing conditions.

To summarise, the demand is subdued and is not healthy (for instance, corporates seek loans for the refinancing of old debts, but not for investment), and supply is tightened, despite it exceeds demand. Nevertheless, there are no doubts that balance sheet clean-up for all sectors is needed. Until such changes take place, despite the stability and high liquidity of the domestic banking sector supported by an expansive monetary policy, the conclusion may be drawn that the scope of monetary policy to encourage credit growth is limited.

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FORECASTING MORTGAGES: INTERNET SEARCH DATA AS A PROXY FOR MORTGAGE CREDIT DEMAND

Branislav Saxa*

Abstract

This paper examines the usefulness of Google Trends data for forecasting mortgage lending in the Czech Republic. While the official monthly statistics on mortgage lending come with a publication lag of one month, the data on how often people search for mortgage-related terms on the internet are available without any lag on a weekly basis. Growth in searches for mortgages and growth in mortgages actually provided are strongly correlated. The lag between these two growth rates is two months. Evaluation of out-of-sample forecasts shows that internet search data improve mortgage lending predictions significantly. In addition to forecasting performance evaluation, an experimental indicator of restrictively tight mortgage credit standards and conditions is proposed. While nowadays many countries run bank lending surveys to monitor the tightness of bank lending standards and conditions, the proposed indicator could be useful in countries without such a survey.

JEL Codes: C22, C82, E27, E51,

Keywords: Google econometrics, Internet search data, Mortgage, Forecasting, Forecast evaluation, Smoothing, Credit demand, Credit supply, Credit standards and conditions

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1. Introduction

In their seminal paper, Choi and Varian (2009a) show how Google Trends data improve near-term forecasts of several economic indicators, including retail sales, car sales, home sales and travel destinations. Other papers followed, showing how unemployment, private consumption and house prices can be forecasted using internet search indices. The official statistics on mortgage lending are published monthly and come typically with a lag of several weeks.[1] In contrast, Google Trends data on the search volumes of queries that users enter into Google are available on a weekly basis without any lag. This paper explores to what extent the Google Trends data can be helpful in predicting mortgage growth.

The analyses in this paper are based on the data for the Czech Republic. Google's share of the search engine market in the country was 71% and internet penetration was 73% in 2013.[2]

Mortgage lending in the Czech Republic has a relatively short history, but the credit market is highly competitive. The volume of new mortgages almost tripled between 2004 and 2007, while real estate prices almost doubled. The demand for mortgages decreased in 2008–2010, but returned to solid growth afterwards. New mortgages doubled between 2010 and 2014, to large extent on the back of increasing importance of mortgage refinancing[3] and decreasing attractiveness of building savings. Volumes of new mortgages have often been influenced by factors other than economic growth, interest rates and inflation, so forecasting models based on these variables struggle to provide reliable forecasts. In this situation, data on how often people google for mortgage rates might contain information useful for improving near-term mortgage growth forecasts.

In the main part of the paper, we assume the supply of mortgages is not limited and we use Google Trends data to forecast mortgage growth. At the end of the paper, it is assumed that the willingness of banks to provide mortgages changes over time and banks might be less willing to lend in certain periods. By comparing the demand for mortgages (proxied by the amount of internet searching for mortgages) and the amount of mortgages actually provided, we construct an experimental indicator which can signal tightening of bank lending standards and conditions. While nowadays many countries run bank lending surveys to monitor the tightness of bank lending standards and conditions, the proposed indicator could be useful in countries without such a survey.

The paper is structured as follows. The next section reviews the economic literature on the use of internet search data for nowcasting and near-term forecasting. Section 3 describes the data employed and provides stylised facts. In section 4, a forecasting exercise shows the usefulness of internet search data for predicting mortgage growth under the assumption that the mortgage supply does not change. Subsequently, an experimental indicator of restrictively tight credit standards and conditions is proposed for cases where this assumption can be dropped and banks are assumed to limit the credit supply in certain periods. The fifth section concludes.

2. Internet search data in the economic literature

The potential of internet search data was first demonstrated in the work of Ginsberg et al. (2009), who suggested a method to analyse Google search queries to track influenza-like illness in a population. The usefulness of internet search data for economic nowcasting and forecasting is demonstrated in Choi and Varian (2009a, 2009b, 2012). In their examples, simple autoregressive models are augmented with search engine data to produce near-term forecasts of automobile sales, unemployment claims, travel destination planning and consumer confidence. In most of their examples, the authors find a reduction in the mean absolute error coming from out-of-sample one-step-ahead forecasting exercises. In the case of initial claims for unemployment benefits, Google Trends data help with the identification of turning points.

[1] The publication lag for mortgage loan statistics in the Czech Republic is one month.

[2] Source: The Webcertain Global Search and Social Report 2013 (http://internationaldigitalhub.com/en/publications/the-webcertain-global-search-and-social-report-2013)

[3] Refinanced mortgages appear as new mortgages in the statistics too.

Several studies showing how internet search data improve predictions followed. Askitas and Zimmerman (2009) perform a forecasting exercise on German unemployment data, showing the potential of internet search data in unemployment predictions. D'Amuri and Marcucci (2012) propose the use of an index of internet job-search intensity as the best leading indicator to predict the US unemployment rate. Fondeur and Karamé (2013) employ the unobserved components approach and use the Kalman filter to estimate a model for nowcasting and forecasting French youth unemployment. Pescyova (2011) uses the data on unemployment in Slovakia, showing that internet search data improve in-sample predictions substantially. Predictions of unemployment in the UK can also be improved using internet search data, as McLaren and Shanbhogue (2011) show. In addition, they illustrate that in the case of house prices, predictions using internet search data can outperform some existing indicators. The extent to which cross-sectional differences in home prices can be predicted using internet search data is studied in Beracha and Wintoki (2013). Schmidt and Vosen (2009) show how Google Trends beats the forecasting performance of the two most common indicators of private consumption in the U.S. (the University of Michigan Consumer Sentiment Index and the Conference Board Consumer Confidence Index).

3. The data and stylised facts

Two time series are used for the analyses in this paper – the nominal volume of new mortgages provided to households by banks in the Czech Republic, and Google Trends data on the search volumes of mortgage-related Czech words with and without diacritics[4] searched for from computers in the Czech Republic.

The statistics on newly provided mortgages are available on a monthly basis and are published by the Czech National Bank one month after the end of the month to which they apply. Data on search volumes of specific terms are available from www.google.com/trends as an index.[5] These data are available on a weekly basis without any publication lag.

For data downloaded at different times, Google Trends returns different data series for the same search query, presumably calculated using different random samples of searches. As the differences between the time series can be rather substantial for the search terms and geographical location used in this paper, ten different data series obtained using the same query at different times were averaged for further use. The pairwise correlations between the ten individual search data series are in the range of 0.78–0.85, while the pairwise correlations between the individual series and the averaged series are in the range of 0.90–0.93. The standard deviations of the individual series range from 11.0 to 12.7, and the standard deviation of the averaged series is 11.0.

Once downloaded and averaged, the search volume data were transformed into monthly frequency in the following way. First, the weekly search volume data were transformed into daily data so that the index for each day was set equal to the value of the index for the relevant week. Subsequently, the daily data were transformed into monthly data so that the monthly value of the index was set to be equal to the average of all the daily values of the index for the respective month.

The sample period used for the analyses shown in this paper is 2007m1–2014m10. Although Google Trends data are available from 2004, the individual search data time series contain many zeros before 2007, presumably due to a low number of related searches (Google Trends sets the index value to 0 if the number of searches does not exceed a certain threshold).

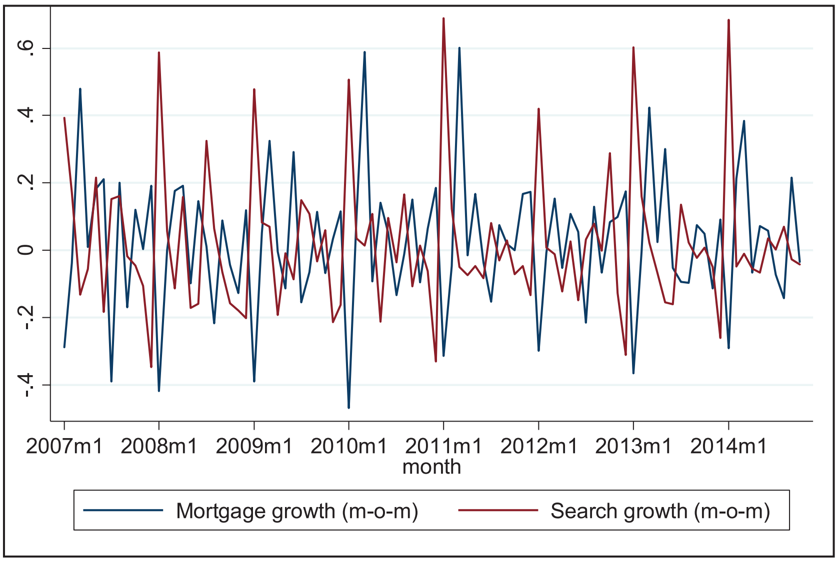
Both time series are shown in levels in Graph A1 in the appendix. For further use, the series are transformed into month-on-month growth rates. These are shown in Graph 1. Both series are stationary;

4 "hypotéka" + "hypoteka" + "hypoteční" + "hypotecni" + "hypotéku" + "hypoteku" + "hypotéky" + "hypoteky" + "úvěr na bydlení" + "uver na bydleni"

5 Google Trends provides search data that are already normalised (divided by a common variable, such as total searches, to cancel out the variable's effect on the data). Values are therefore between 0 and 100.

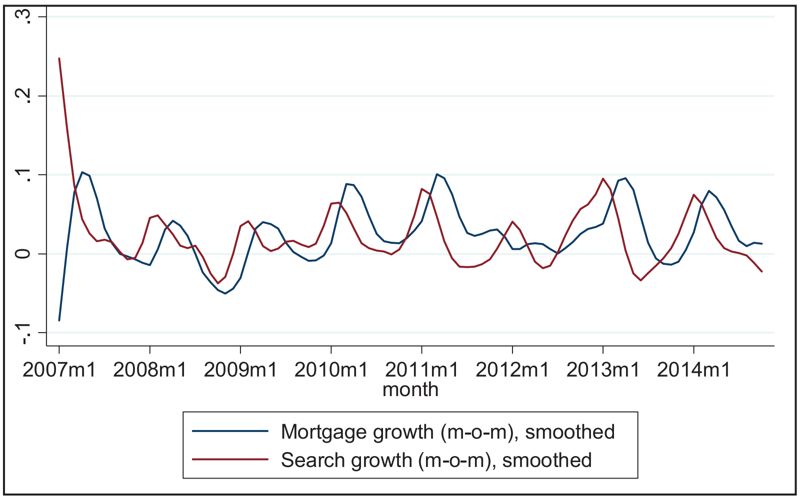
the augmented Dickey–Fuller test strongly rejects the null hypothesis of a unit root in both cases. The two series are highly seasonal.[6]

Graph 1: Month-on-month growth rates of mortgages and searches



To inspect for potential co-movement of the two growth rates, several smoothing methods were applied. Graph 2 shows the growth rates smoothed using the Hodrick–Prescott filter with lambda set to 10.[7] While the beginning and the end of sample suffer from the typical Hodrick–Prescott end-point bias, the search growth rate in general mimics the mortgage growth rate reasonably well, with an obvious lead. The cross-correlogram for the whole sample period (Graph A3 in the appendix) suggests that searches lead mortgages by two to three months. The correlations of the growth rates at these lags are 0.75 and 0.74 respectively.

Graph 2: Smoothed (HP filtered) month-on-month growth rates of mortgages and searches



However, just by looking at Graph 2, one cannot reject the hypothesis that the lag between searches and mortgages changes over time. When cross-correlations are calculated for three different overlapping subsamples (2007m1–2009m12, 2009m1–2011m12, 2011m1– 2014m10), the correlation

[6] Google Trends data also have a specific type of seasonality. Search volume indices tend to decrease towards the end of the year as a result of the overall volume of searches being inflated by Christmas-related shopping searches.

[7] The lambda parameter is substantially lower than the values typically used for monthly frequency data. This is because the aim here is to obtain smoothed short-term changes, not to isolate business cycle frequencies from the long-run trend component. The result of alternative smoothing using a symmetric moving-average filter with $\hat{x} = \frac{1}{9}(1x_{t-2} + 2x_{t-1} + 3x_t + 2x_{t+1} + 1x_{t+2})$ is depicted in Graph A2 in the appendix (the search growth rate is lagged by two months in Graph A2).

is strongest at the three-month lag in the first period, but at the two-month lag in the second and third periods (details in Table A1 in the appendix).

To see whether the lag structure changes over time, rolling window correlations between the smoothed growth rates of mortgages and searches are calculated for lags 0 to 4. The window width is 48 months. Graph A4 in the appendix shows how the correlations at different lags change over time (the times on the horizontal axis indicate the end of the subsample used for the calculation of the correlation). The correlation is strongest for the two-month lag over the whole sample period, but the correlation at the three-month lag gets equally high towards the end of the sample.

Table A2 in the appendix shows summary statistics for all the variables used.

4. Empirical approach and results

4.1. Forecasting mortgages

In the first part of this paper, Google Trends data are used to forecast mortgage lending. A simple autoregression process with and without a seasonal component is estimated to judge whether internet search volume data can improve the mortgage lending forecast.

To see how much of the monthly dynamics in mortgage lending can be captured by the amount of googling for mortgages two months earlier, the month-on-month mortgage growth is first regressed on its lagged values (referred to as the AR(1) model). The results are then compared to the results of the same regression augmented with the month-on-month growth in searching for mortgages, lagged by two months (referred to as the ARX model). As Table 1 shows, the amount of variation explained by the regression (proxied by adjusted R-squared) increases substantially, from 0.05 to 0.39.

Table 1: Variation in mortgage lending explained by amount of searching two months earlier
(least squares estimation; the dependent variable is month-on-month growth of mortgage lending; standard errors in parentheses)

	AR(1)		ARX	
L.Mortgage growth (m-o-m)	-0.24	**	-0.41	***
	(0.10)		(0.09)	
L2.Search growth (m-o-m)			0.58	***
			(0.08)	
Constant	0.03	*	0.03	
	(0.02)		(0.02)	
Adjusted R-squared	0.05		0.39	
Number of observations	93		92	

*Note: * p<0.10, ** p<0.05, *** p<0.01*

In the next step, the usefulness of search growth data in forecasting mortgage lending is assessed in an out-of-sample forecasting exercise.[8] The baseline forecast of mortgage growth is compared with the forecast exploiting the data on search growth lagged by two months. The two estimated equations take the following form:

$$\text{AR(1)}: \Delta mortgage_t = \alpha + \beta \Delta mortgage_{t-1}$$

and

$$\text{ARX}: \Delta mortgage_t = \alpha + \beta \Delta mortgage_{t-1} + \gamma \Delta search_{t-2}$$

8 Performed in STATA using the STATICFC module authored by Baum (2013).

The window used for the estimation of the equations' coefficients extends from 2007m1– 2008m8 to 2007m1–2014m9.[9] The one-month and two-months-ahead forecasts constructed using the estimated coefficients are then compared with the actually observed values and the forecast errors are calculated. Two forecast evaluation measures are shown in Table 2: the mean absolute error (MAE) and the square root of the mean-squared forecast error (RMSE). Inclusion of lagged search growth reduces the MAE and RMSE of the one-step-ahead mortgage forecasts by approximately 18% and 23% respectively. A similar improvement is observed with the two-steps-ahead forecasts.[10] In both cases, the Diebold–Mariano test strongly rejects the null hypothesis that the forecast accuracy of the models with and without lagged searches is the same. Graphs A5 and A6 in the appendix show the observed mortgage growth as well as the one-step-ahead predictions without and with search growth used as an explanatory variable.

Table 2: Mortgage growth out-of-sample forecasting exercise

(Mean absolute errors (MAE) and square roots of mean-squared forecast error (RMSE) for one-month and two-months-ahead forecasts)

	AR(1)	ARX	Change	Diebold-Mariano	
				S(1)	p-value
One-step-ahead forecast					
MAE	0.1411	0.1162	-18%		
RMSE	0.1919	0.1475	-23%		
				4.25	0.00
Two-steps-ahead forecast					
MAE	0.1420	0.1150	-19%		
RMSE	0.1924	0.1466	-24%		
				4.27	0.00

Finally, mortgage growth lagged by 12 months is added to the regressions so that all seasonal effects can be captured by this term. Most of the variation in mortgage growth is now explained by seasonality. Nevertheless, lagged search growth still improves the fit, as Table 3 shows.

Table 3: Variation in mortgage lending explained by amount of searching two months earlier and seasonal term

(least squares estimation; the dependent variable is month-on-month growth of mortgage lending; standard errors in parentheses)

	SAR(1)	SARX
L.Mortgage growth (m-o-m)	-0.17 **	-0.28 ***
	(0.08)	(0.07)
L12.Mortgage growth (m-o-m)	0.67 ***	0.47 ***
	(0.07)	(0.08)
L2.Search growth (m-o-m)		0.35 ***
		(0.08)
Constant	0.01	0.01
	(0.02)	(0.01)
Adjusted R-squared	0.53	0.61
Number of observations	82	82

*Note: * p<0.10, ** p<0.05, *** p<0.01*

9 2007m1–2008m8 to 2007m1–2014m8 in the case of the two-months-ahead forecast comparison

10 A number of alternative specifications were estimated too. Two versions with smoothed lagged search growth rates (HP and moving average) in the ARX equation beat the forecasting performance of the AR(1) model, but their performance is worse than that of the ARX model with non-smoothed lagged search growth rates. A version with non-smoothed lagged search growth rates entering the ARX equation with a lag of three months instead of two months does not beat the forecasting performance of the AR(1) model.

The two equations estimated in the out-of-sample forecasting exercise now look as follows:

SAR(1): $\Delta mortgage_t = \alpha + \beta\Delta mortgage_{t-1} + \theta\ \Delta mortgage_{t-12}$

and

SARX: $\Delta mortgage_t = \alpha + \beta\Delta mortgage_{t-1} + \theta\ \Delta mortgage_{t-12} + \gamma\Delta search_{t-2}$

The results of the out-of-sample forecasting exercise with a seasonal term are summarised in Table 4. The inclusion of lagged search growth reduces the MAE and RMSE of the one-stepahead mortgage forecasts by approximately 8% and 10% respectively. The reductions in the forecast error for the two-steps-ahead forecasts are approximately 7% and 10% respectively. However, as the differences in forecasting performance between SAR(1) and SARX are smaller, the Diebold–Mariano test does not reject the null hypothesis that the forecast accuracies of the two competing models are equal at any conventional significance level (the p-values are 0.13 and 0.16 for the one-step and two-steps-ahead forecasts respectively).

Table 4: Mortgage growth out-of-sample forecasting exercise with seasonal term

(Mean absolute errors (MAE) and square roots of mean-squared forecast error (RMSE) for one-month and two-months-ahead forecasts)

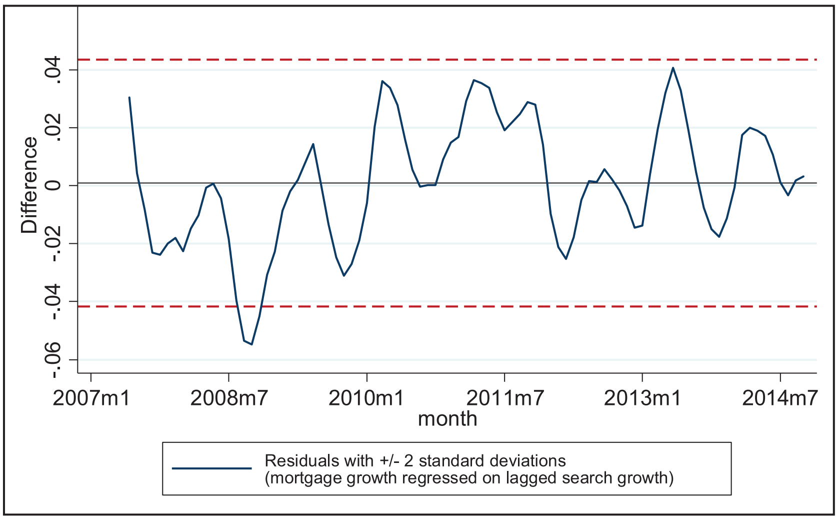
	SAR(1)	SARX	Change	Diebold-Mariano	
				S(1)	p-value
One-step-ahead forecast					
MAE	0.0985	0.0909	-8%		
RMSE	0.1299	0.1168	-10%	1.50	0.13
Two-steps-ahead forecast					
MAE	0.0992	0.0925	-7%		
RMSE	0.1307	0.1182	-10%	1.41	0.16

4.2. Experimental indicator of restrictively tight bank lending standards and conditions

In the second part of the paper, demand for mortgages is compared with mortgages actually provided and an experimental indicator of restrictively tight bank lending standards and conditions is constructed. In this part, it is assumed that the willingness of banks to provide mortgages changes over time and in some periods fewer mortgages are provided not due to lower demand, but because of restricted supply. While nowadays many countries run bank lending surveys to monitor the tightness of bank lending standards and conditions, the proposed indicator could be useful in countries without such a survey.

The idea behind the indicator is straightforward. The smoothed growth rate of mortgages actually provided is regressed on the smoothed growth rate of searches lagged by two months. The residuals from this regression represent the part of the variation in mortgages that cannot be explained by the variation in demand for mortgages. Growth of demand substantially above the growth of mortgages actually provided can signal a lower willingness of banks to provide mortgages.

Graph 3: Experimental indicator of restrictively tight bank lending standards and conditions[11]



If the lagged amount of searching grows faster than the amount of mortgages, the indicator is negative and suggests that fewer mortgages are provided than demanded. Graph 3 shows the values of the experimental indicator for the Czech Republic, along with the two-standarddeviation band around the mean of the indicator.[12] The indicator leaves the band only in the third quarter of 2008.

What happened with credit standards and conditions in this period (which saw the outbreak of the financial crisis)? The bank lending survey did not yet exist in the Czech Republic in 2008, but it is possible to check the bank lending survey for the Eurozone. The net tightening of credit standards applied to loans to households for house purchase reached 36% in the Eurozone in the third quarter of 2008. This is the second-highest number in the history of the Eurozone bank lending survey.[13] The period identified using the suggested experimental indicator thus very likely coincides with the period of most significant tightening of credit standards and conditions during the period analysed.

Of course, reasons other than excessively tight credit standards and conditions might influence the value of the experimental indicator too. These could include changing behaviour of people searching for mortgages and a changing lag between searching for a mortgage and signing a mortgage contract.

Two alternative specifications were considered. If a 1:1 relationship is imposed on the growth rates of mortgages and lagged searches, the indicator can be constructed as a simple difference of growth rates. As Graph A9 in the appendix shows, this approach delivers qualitatively similar results. Finally, if a lag of three months is assumed instead of a lag of two months, the extreme values of the indicator are different but the dynamics of the indicator remain the same (Graph A10 in the appendix).

[11] As the smoothed series suffer from the typical Hodrick–Prescott end-point bias and the indicator becomes very low in the first few months, the chart begins in 2007m6.

[12] Under the assumption that the residuals are distributed normally, the indicator leaves the two-standarddeviation band in approximately 5% of cases.

[13] The only higher number was reported one quarter later.

5. Conclusion

This paper investigates the usefulness of internet search data for forecasting mortgage lending. While the official monthly statistics on mortgage lending come with a publication lag of one month, the data on how often people search for mortgages on the internet are available without any lag on a weekly basis. As this paper shows, the growth rates of searches and mortgages are strongly correlated and the volume of searches leads the volume of mortgages provided by two months. The variation in month-on-month search growth explains a substantial part of the variation in month-on-month mortgage growth. Most importantly, outof-sample near-term forecast exercises show that the volume of searches improves the short-term predictions of mortgage lending.

When predicting mortgages using internet search data, it is assumed that the supply of mortgages is unrestricted and the amount of mortgages provided is equal to the demand for mortgages. In the last part of the paper, an experimental indicator of restrictively tight mortgage credit standards and conditions is proposed. Once the assumption of unrestricted credit supply is dropped, the willingness of banks to provide mortgages can change over time. In certain periods, fewer mortgages are provided not due to lower demand, but because of restricted supply. The proposed indicator identifies the third quarter of 2008, i.e. the outbreak of the financial crisis, as a period of substantially restricted credit supply. This is in line with information on credit standards and conditions from the bank lending survey for Eurozone countries. While nowadays many countries run bank lending surveys to monitor the tightness of bank lending standards and conditions, the proposed indicator could be useful in countries without such a survey.

Together with the available studies on forecasting unemployment and private consumption, the two applications in this paper illustrate the usefulness of internet search information for monetary policy and economic forecasting in general.

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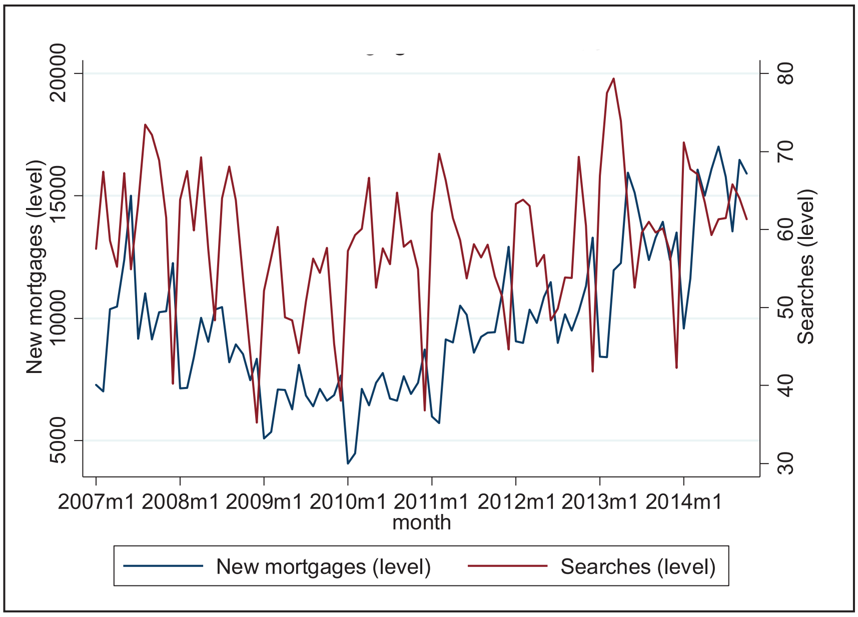
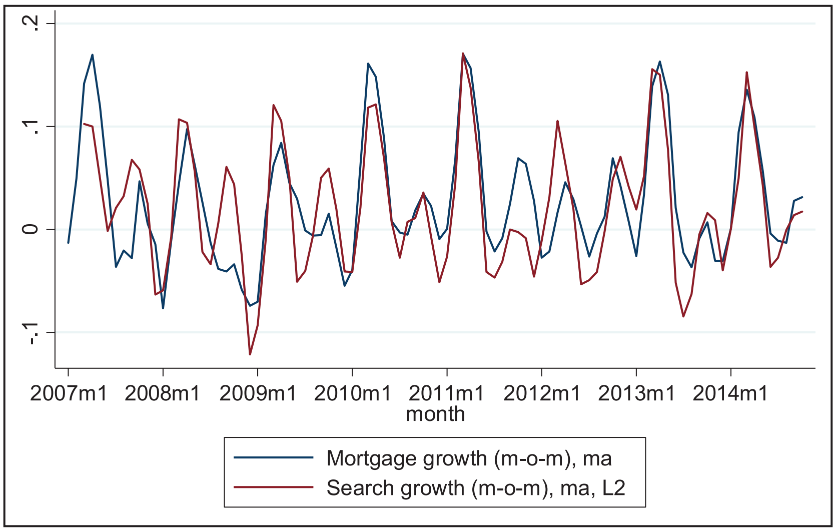
Appendix

Graph A1: Levels of mortgages and searches



Graph A2: Smoothed month-on-month growth rates of mortgages and searches (moving average)



Graph A3: Cross-correlogram between smoothed m-o-m growth rates of mortgages and searches

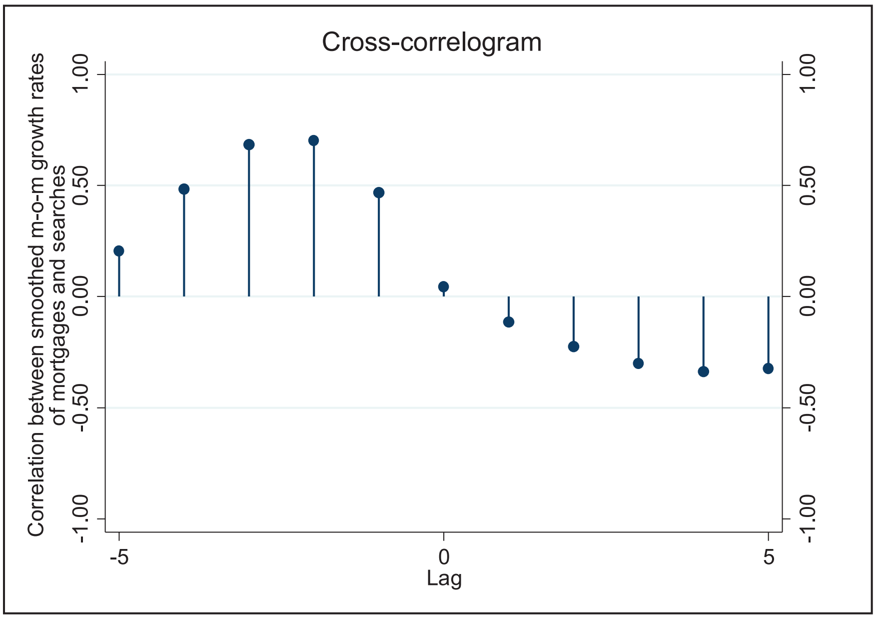


Table A1: Cross-correlations between mortgages and searches for different lags and subsamples
(significance levels in parentheses, lags with the highest correlation coefficient in bold)

Lag in months	Subsample			Whole sample
	2007m1-2009m12	2009m1-2011m12	2011m1-2014m10	2007m1-2014m10
0	-0.08	0.20	0.24	0.04
	(0.66)	(0.24)	(0.11)	(0.67)
1	0.47	0.62	0.60	0.49
	(0.00)	(0.00)	(0.00)	(0.00)
2	0.78	**0.83**	**0.84**	**0.75**
	(0.00)	**(0.00)**	**(0.00)**	**(0.00)**
3	**0.81**	0.74	0.83	0.74
	(0.00)	(0.00)	(0.00)	(0.00)
4	0.67	0.45	0.63	0.54
	(0.00)	(0.01)	(0.00)	(0.00)

Graph A4: Rolling window correlations between smoothed m-o-m growth rates of mortgages and searches at different lags

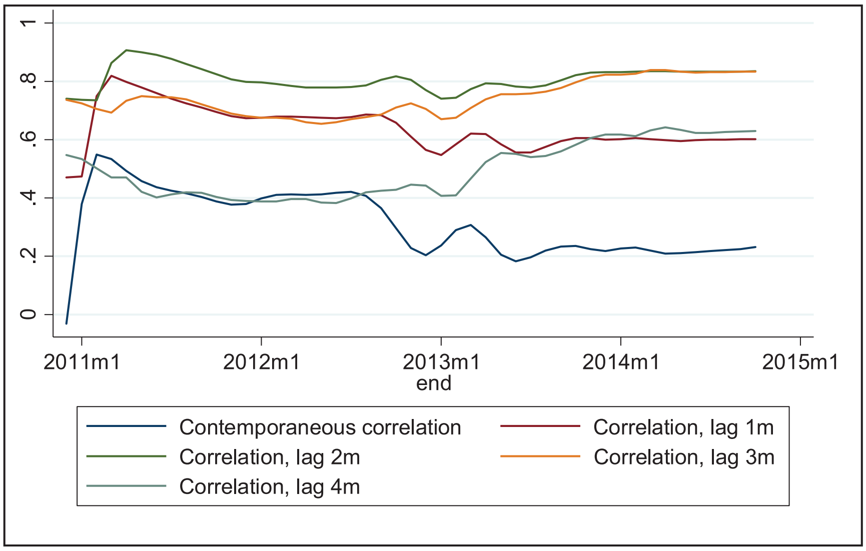
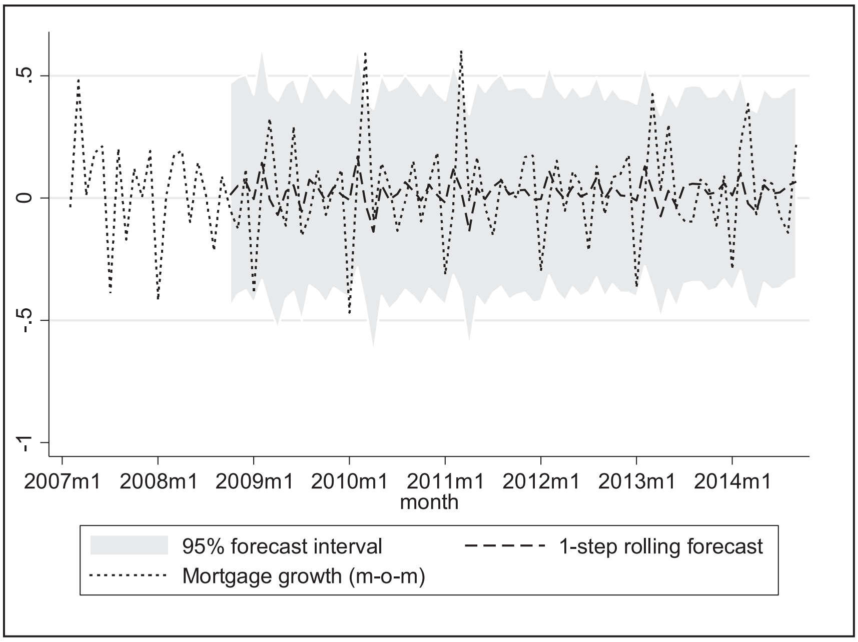


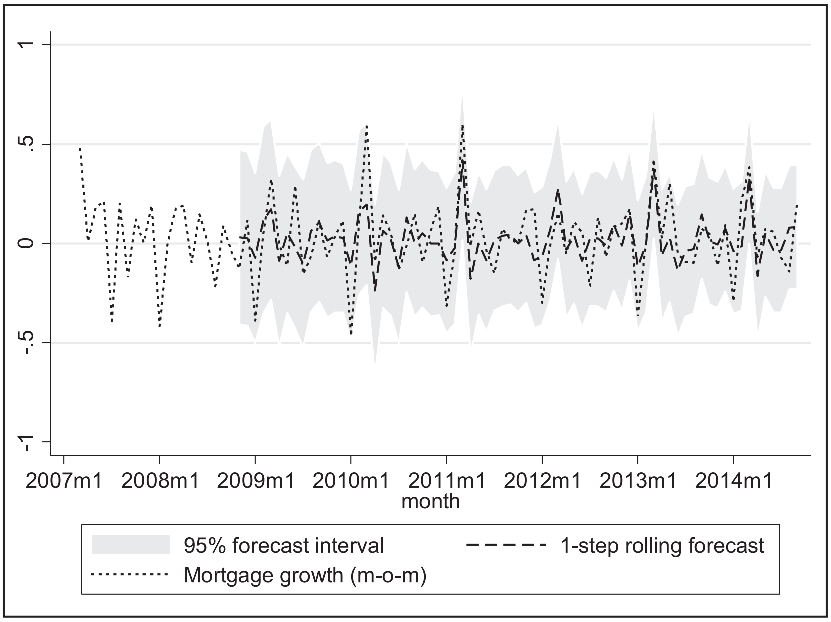
Table A2: Summary statistics

Variable	**Obs**	**Mean**	**Std. Dev.**	**Min**	**Max**
mortgages	94	9831.5	3027.5	4074.5	17021.4
searches	94	58.5	8.8	35.2	79.4
m-o-m mortgage growth	94	0.0250	0.2000	-0.4675	0.6015
m-o-m search growth	94	0.0224	0.2053	-0.3468	0.6889
smoothed m-o-m mortgage growth (HP filter, λ=10)	94	0.0250	0.0367	-0.0842	0.1031
smoothed m-o-m search growth (HP filter, λ=10)	94	0.0224	0.0406	-0.0376	0.2477
experimental index	92	0.0000	0.0234	-0.0943	0.0407

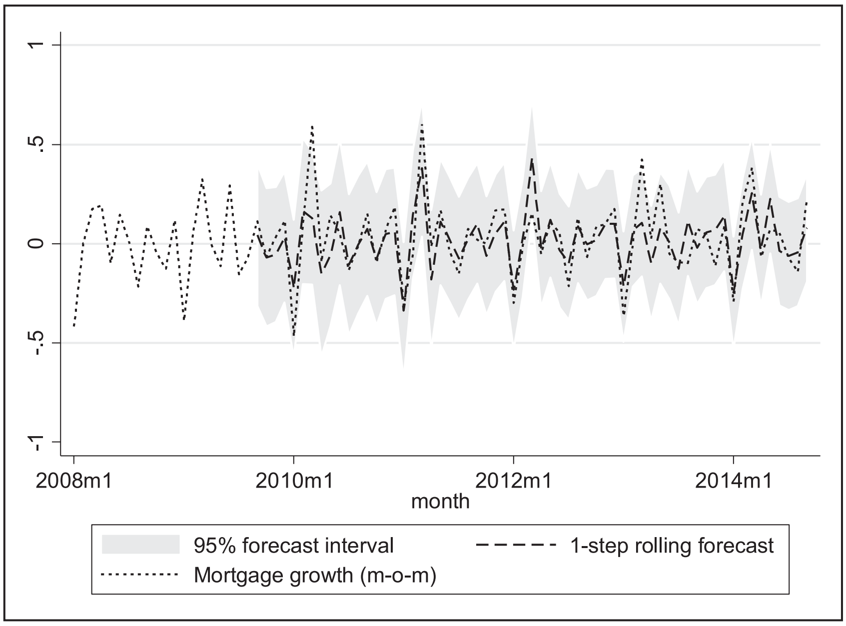
Graph A5: One-step-ahead out-of-sample forecasts of month-on-month growth rate of mortgages (without seasonal term, without search growth)



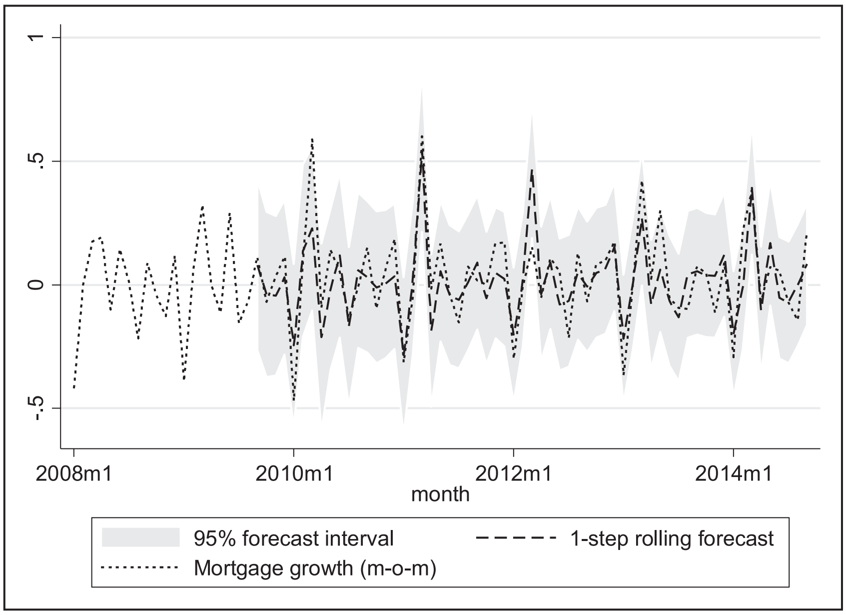
Graph A6: One-step-ahead out-of-sample forecasts of month-on-month growth rate of mortgages (without seasonal term, with search growth)



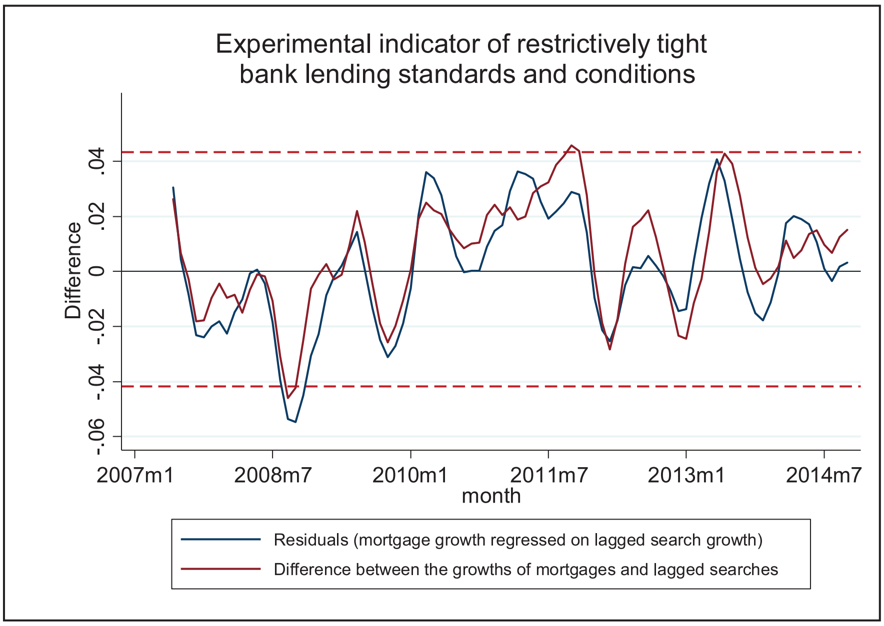
Graph A7: One-step-ahead out-of-sample forecasts of month-on-month growth rate of mortgages (with seasonal term, without search growth)



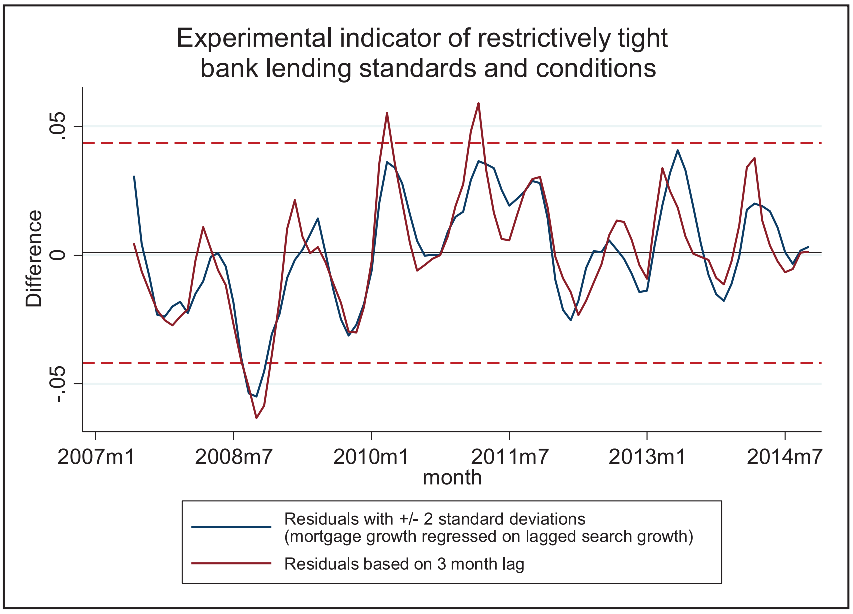
Graph A8: One-step-ahead out-of-sample forecasts of month-on-month growth rate of mortgages (with seasonal term, with search growth)



Graph A9: Comparison of baseline experimental indicator with version constructed as simple difference of growth rates.



Graph A10: Comparison of baseline experimental indicator with version assuming lag of three months instead of two months



NATIONAL BANK OF THE REPUBLIC OF MACEDONIA

P R O G R A M

4th Research Conference “Addressing Structural Rigidities in View of Monetary Policy Transmission Effectiveness”
23 April 2015, Skopje

23 April 2015 (Thursday)

9.00 Dimitar Bogov, Governor of the National Bank of the Republic of Macedonia, Opening speech

<u>9.10-11.00 Session I - Keynote lecture and high level policy discussion panel</u>

9.10 Miroslav Singer, Governor of the Czech National Bank, Keynote lecture

Chair: Aleksandar Stojkov, Council member, NBRM

9.40 Boris Vujčić, Governor of the Croatian National Bank, Discussant

9.50 Panel discussion: Miroslav Singer, Boris Vujčić, Dimitar Bogov

10.40 Questions from the audience

10.55 Announcement of the Annual Award of the NBRM for the best paper in macroeconomics and banking by a young researcher

11.00 Coffee break

<u>11.30 - 13.00 Session II: Structural rigidities, growth and monetary policy</u>

Chair: Vladimir Filipovski, Faculty of Economics, Skopje

11.30 Balázs Égert, OECD, Structural Policies and Economic Growth: the Impact Product and Labour Market Policies on MFP, Investment and Labour Market Outcomes

11.50 Hülya Saygili, Central Bank of the Republic of Turkey, Trade in Goods, Globalization in Production Structure and Inflationary Dynamics: Cross Country Evidence

12.10 Magdalena Petrovska, NBRM, Florian Huber, OeNB, Price and Wage Rigidities in the Republic of Macedonia: Survey Evidence from Micro-Level Data

12.30 Altin Tanku, Bank of Albania, Discussant

12.45 Discussion

13.00 **Lunch**

14.00 - 15.30 Session III: The changing nature of the monetary policy transmission mechanism

Chair: Ana Mitreska, Director of the Monetary Policy and Research Department, NBRM

14.00 Yannick Lucotte, ESG Management School, France, Mr. Grégory Levieuge, Mr. Sebastien Ringuedé, Central Bank Credibility and the Expectation Channel: Evidence Based on a New Credibility Index

14.20 Utku Özmen, Mr. Çağrı Sarikaya, Central Bank of the Republic of Turkey, Sensitivity of Inflation to Demand Conditions in Turkey: Determining CPI Items Responding to Output Gap and Credits

14.40 Lenarčič Črt, Bank of Slovenia, Is There a Harrod-Balassa-Samuelson Effect Present in the Data? New Quarterly Panel Data Evidence from 25 European Countries

15.00 Marjan Petreski, University American College Skopje, Discussant

15.15 Discussion

15.30 Coffee break

15.45 - 17.00 Session IV: Structural features of the financial system

Chair: Aneta Krstevska, Chief Economist, NBRM

15.45 Mirna Dumičić, Igor Ljubaj , Croatian National Bank, Forensic Analysis of Credit Activity in Croatia

16.05 Branislav Saxa, Czech National Bank, Forecasting Mortgages: Internet Search Data as a Proxy for Mortgage Credit Demand

16.25 Alessio Ciarlone, European Central Bank, Discussant

16.45 Discussion

17.00 Wrap up and closing of the conference

* * *