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Effective monetary policy conservatism: A comparison of 11 OECD countries

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Abstract

Modern monetary economists argue that institutional aspects such as central bank independence and central bank conservatism play an important role for the performance of an economy. In order to be able to compare the effects of different institutions it is necessary to measure both central bank independence and conservatism. In this paper we propose a new methodology of uncovering the degree of effective monetary policy conservatism from observed central bank behavior. Employing a variant of the Barro-Gordon-model we derive an optimal prime rate reaction function and show that more effectively conservative monetary policy tends to react less active to shocks to the real economy. In order to illustrate the proposed methodology we then estimate a common prime rate reaction function for a sample of 11 central banks in a panel setting and allow the reaction to real disturbances to differ between countries.

JEL classification: E31; E58

Keywords: central banking, conservatism, central bank independence, inflation

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1 Introduction

Until the early 1980s monetary policy was typically modeled as a control problem. In the models employed central banks make use of their instruments in order to stabilize the economy or, if possible, to contribute to a higher real growth rate, thereby treating the macroeconomic environment as given. Moreover, the design of monetary policy institutions was almost completely neglected. This view changed in the aftermath of Lucas' (1976) critique of the prevailing macroeconomic policy evaluation methods, which predicted the effects of policy experiments purely on correlations in historical data. The insight of the necessity to predict how market participants will react to policy changes led to the development of the game theoretic approach of monetary policy. Building up on earlier work by Kydland and Prescott (1977) on time-inconsistent policies, Barro and Gordon (1983) argue that monetary policy can be interpreted as a game between the monetary authority and wage bargainers. The basic message of the model is that a monetary authority caring about stable prices yet also having an ambitious output target will cause excessive inflation without any real effects. Moreover, the model predicts the inflation bias to be the larger, the more relative weight is attached to the ambitious output (or employment) target.

In the aftermath of the publication of the Barro-Gordon-model several authors expressed their doubts on the Barro-Gordon-model as an empirical relevant positive theory of inflation.[1] To some extent, these doubts evolved as a consequence of the fact that for a long period of time no serious attempts were made at testing for

[1] See e.g. Taylor (1983), Blinder (1997) and Forder (2004). An interesting review article on this discussion was provided by Blackburn and Christensen (1989).

inflation biases (Walsh 2003, p. 412). Nevertheless, well before Romer (1993), Ireland (1999) and Berlemann (2005) presented empirical evidence in favor of the Barro-Gordon-model, the model established quite quickly as a standard framework to study the behavior of monetary authorities.

The Barro-Gordon-model inspired a large literature on how to overcome the time inconsistency problem of monetary policy.[2] Among the most prominent proposals is the delegation solution, suggesting the government hand over the sole responsibility of monetary policy to an independent central bank. However, central bank independence is only a necessary but not a sufficient condition to ease the inflation bias. This is due to the fact that delegating monetary policy to an independent central bank sharing the same objective function as the government would not lead to a reduction of the inflation bias. In order to mitigate the inflation bias a somewhat independent central bank has to be governed by a "conservative" central banker, putting less relative weight on the ambitious output target than the government (and thus the median voter). In an influential article, Rogoff (1985) showed how to detect the optimal degree of conservatism of a central banker for the case of a completely independent central bank. He thereby contributed considerably to deepen our understanding of the importance of institutional aspects for monetary policy.

Maybe the most prominent example of a country with an an independent central bank and conservative central bank is Germany. Before the German Bundesbank became part of the European System of Central Banks it ranked among the most independent central banks of the world, regardless of the applied ranking method.[3] Moreover, the presidents of the German Bundesbank had a strong reputation for being highly inflation averse. Of course, the German solution was chosen well

[2] See Svensson (1997) or Walsh (2003) for a review of the related literature.

[3] See e.g. Grilli et al. (1991) or Cukierman (1992). A review and critical assessment of existing measures can be found in Forder (1999).

before the academic debate on the importance of central bank independence and conservatism unfolded. However, the academic debate provided the economic rationale for this concept and was surely one of the driving forces of the observable international trend towards increasing the degree of central bank independence.[4]

Inspired by Rogoff's work and the growing interest of monetary authorities a substantial literature on the empirical measurement and the effects of central bank independence evolved. Various measures of legal and factual central bank independence were developed and related to macroeconomic variables.[5] However, up to now comparatively few attempts were undertaken to study the degree of conservatism. In the light of the fact that central bank independence without conservatism is insufficient to solve the problem of time inconsistency, this is a severe shortcoming.[6] Empirical evidence on the consequences of central bank independence might be distorted when differing levels of conservatism are not taken into account. A more meaningful measure would be a joint measure of (i) central bank independence and (ii) the central bank's degree of conservatism and (iii) the government's (and thus the median voter's) degree of conservatism, i.e. a measure of "effective monetary policy conservatism".

This paper fills the described gap in the literature by proposing a simple method of measuring the degree of average effective monetary policy conservatism of modern central banks. The basic idea is to uncover effective monetary policy conservatism on the basis of observed central bank behavior. Employing a variant of the Barro-Gordon-model we derive an optimal prime rate reaction function and show that more effective conservative monetary policy tends to react less active

[4] See e.g. Berger et al. (2001) or Eijffinger et al. (1996). Forder (2005) offers a variety of explanations of the spread of central bank independence.

[5] A survey of this literature can be found in Berger et al. (2001).

[6] Furthermore, the literature often fails to distinguish properly between the concepts of central bank independence and conservatism. For a similar view see Lippi (2000) or Berger and Woitek (2005).

to shocks to the real economy.[7] In order to illustrate the proposed methodology we then estimate a common prime rate reaction function for a sample of 11 central banks in a panel setting and allow the reaction to real disturbances to differ between countries.

The paper is organized as follows: In section 2 we introduce the theoretical framework and derive a central bank's optimal prime rate reaction function. Section 3 deals with a brief review of the empirical literature on central bank independence and conservatism. In section 4 we present our empirical design, the data and the estimation results. The paper ends with a summary of the main findings and some conclusions.

2 Effective conservatism in a game-theoretic model of monetary policy

Our theoretical model stands in the tradition of the literature which initiated the discussion on central bank independence and conservatism evolved, i.e. the game-theoretic model of monetary policy as proposed by Barro and Gordon (1983) and further developed by Rogoff (1985) and Eijffinger and Hoeberichts (1998). There are two actors in the standard monetary policy game: a monetary authority controlling inflation and trade unions bargaining wages collectively, thereby forming rational expectations on the inflation rate. In line with this literature we employ an expectations-augmented Phillips-curve

$$u_t = \bar{u} + \pi_t^e - \pi_t + \upsilon_t, \tag{1}$$

where u_t denotes (the log of) unemployment in period t, $\bar{u}$ the (log of) natural rate of unemployment, π_t inflation, π_t^e the trade unions' inflation expectations and υ_t a

[7] As Berger and Woitek (2005) state, the implications of effective conservatism with respect to the reaction to macroeconomic shocks depend on structural and dynamic characteristics of the economy. Since the later reported empirical results fit the common perception of effective conservatism quite well, the use of the employed theoretical framework seems to be justified.

white-noise random shock. The standard model of Barro and Gordon (1983) implies that the monetary authority directly and perfectly controls inflation. Following Ruge-Murcia (2003) and Walsh (2003) we assume that the monetary authority imperfectly controls inflation using a policy instrument i which is linked to inflation according to

$$\Delta\pi_t = -\Delta i_t + \eta_t \tag{2}$$

with η being a white-noise control error.

Following Eijffinger and Hoeberichts (1998) the objective function underlying monetary policy might be described by an independence-weighted average of the loss function of the government and the central banker. Defining I as the degree of central bank independence, the monetary authority's loss function can then be denoted as

$$l_t^M = I \cdot l_t^C + (1 - I) \cdot l_t^G. \tag{3}$$

The government's one-period loss function is given by

$$l_t^G = \frac{1}{2} \cdot (\pi_t - \pi^*)^2 + \frac{1}{2} \cdot \beta^G \cdot (u_t - u^*)^2, \tag{4}$$

where β^G is the weight the government assigns to the goal of high employment relative to the goal of stable prices. Following Barro and Gordon (1983) we assume that the government's loss function coincides with the social loss function.

The loss function of a weight-conservative central banker can be denoted as

$$l_t^C = \frac{1}{2} \cdot (\pi_t - \pi^*)^2 + \frac{1}{2} \cdot \beta^C \cdot (u_t - u^*)^2 \tag{5}$$

with $\beta^C = \beta^G - \bar{\beta}$ and $\bar{\beta} > 0$. Inserting (4) and (5) in (3) yields the objective function of the monetary authority

$$l_t^M = \frac{1}{2} \cdot (\pi_t - \pi^*)^2 + \frac{1}{2} \cdot \underbrace{(\beta^G - \bar{\beta} \cdot I)}_{\beta^M} \cdot (u_t - u^*)^2, \tag{6}$$

where $\beta^M = \beta^G - \bar{\beta} \cdot I$ can be interpreted as the degree of effective monetary policy conservatism depending on the degrees of conservatism of the government and the central banker as well as the degree of central bank independence.[8]

The sequential structure of the game is as follows: First, the monetary authority announces an inflation rate. Then, trade unions form their inflation expectations and anticipate them during wage negotiations. Afterwards the shock υ_t realizes and finally the monetary authority determines the rate of inflation and sets its policy instrument i in order to minimize its expected loss.

Formally, the Nash-solution can be calculated by differentiating the monetary authority's loss function with respect to the inflation rate. Taking the rational expectation of the first-order condition, solving for expected inflation and inserting the result into the first-order condition yields

$$\pi_t = \pi^* + (\beta^G - \bar{\beta} \cdot I) \cdot (\bar{u} - u^*) + \frac{\beta^G - \bar{\beta} \cdot I}{1 + \beta^G - \bar{\beta} \cdot I} \cdot \upsilon_t. \tag{7}$$

Thus, we can conclude that the well-known inflation bias depends inversely on the degree of effective monetary policy conservatism ($\beta^M = \beta^G - \bar{\beta} \cdot I$). Obviously, central bank independence and the conservatism of both the government and the central banker are (imperfect) substitutes in easing the inflation bias in the described framework.[9]

With i being the policy interest rate of a monetary authority the optimal interest rate change $\Delta^* i_t$ is calculated by inserting (7) in (2):

$$\Delta^* i_t = -\beta^M \cdot (\bar{u} - u^*) - \pi^* - \frac{\beta^M}{1 + \beta^M} \cdot \upsilon_t + \pi_{t-1} + \eta_t. \tag{8}$$

[8] Note, that lower values of $\beta^G - \bar{\beta} \cdot I$ correspond to a higher degree of effective conservatism.

[9] This result does not hold in the framework used by Lippi (2000). However, his interpretation is quite different from Rogoff's (1985) original line of argument. He defines conservatism as the difference between the median inflation target of society and that of the central bank and thus quite differently from the rest of the literature. In this paper we stick to the more common interpretation of conservatism which is concerned with the weighting parameters.

The stochastic shocks υ_t on unemployment cause fluctuations around the natural level. According to equation (8) central banks will partially offset supply shocks via interest rate variations. For a given unemployment shock υ_t the interest rate variation decreases in effective conservatism.

In equilibrium the shock can be expressed in terms of fluctuations of the rate of unemployment around its natural level, i.e. the unemployment gap $\widehat{u}_t$. Inserting equilibrium inflation and inflation expectations in (1) yields

$$\upsilon_t = (1 + \beta^G - \bar{\beta} \cdot I) \cdot \underbrace{(u_t - \bar{u})}_{\widehat{u}_t}. \tag{9}$$

Hence, we can substitute υ_t in (8) by equation (9):

$$\Delta^* i_t = -\beta^M \cdot (\bar{u} - u^*) - \pi^* - \beta^M \cdot \hat{u}_t + \pi_{t-1} + \eta_t. \tag{10}$$

Thus, the relation between the unemployment gap and the optimal prime rate change reveals information on the central bank's effective degree of conservatism.

3 Previous empirical evidence

Most empirical studies in the field engage in attempts to relate policy outcomes such as inflation or unemployment to different measures of central bank independence or conservatism. Most of these studies find a significant negative correlation between central bank independence and/or conservatism on the one hand and inflation and/or inflation variance on the other.

A first strand of the literature engages in developing indices of statutory central bank independence. In order to do so, authors like Bade and Parkin (1988), Alesina (1989), Grilli et al. (1991), Eijffinger and Schaling (1992), Cukierman (1992), Alesina and Summers (1993), Barro (1995), Loungani and Sheets (1997), Lybek (1999), Maliszewski (2000) and Freytag (2002) analyze the existing legal rules defining the degree of formal central bank independence.[10] While these studies differ considerably in sample countries, sample periods, applied categories and weighting methods almost all of them find significantly negative correlations between statutory measures of central bank independence on the one hand and inflation (and sometimes inflation variance) on the other.[11]

However, especially in developing and transition countries the legal rules differ heavily from factual practices (Forder 1996, 1998). While anecdotal evidence on this hypothesis is available for various countries (Hochreiter and Kowalski 2000), the case of Bulgaria throughout the 1990s provides an excellent example for the differences between de-jure-rules and de-facto-behavior. As described in Berlemann and Nenovsky (2004) the Bulgarian National Bank contributed significantly to excessive inflation and the Bulgarian Financial Crisis of 1996/1997 by continuously monetizing public debt although formally being absolutely independent from the government. As a consequence, adequate measures of central bank independence should be based on informal rules and practices rather than solely on legal codes. This critique led to a second strand of the literature, which is concerned with developing indices of actual central bank independence. For example, Cukierman (1992) developed two proxies of actual independence, the turnover rate of central bank presidents and a survey indicator.[12] Furthermore, Cukierman and Webb (1995) construct a political vulnerability index of central banks for the period from 1950 to 1989. The empirical evidence Cukierman (1992) presents points into the direction that the degree of factual central bank independence is negatively correlated with inflation.

[10]Moreover, there is a considerable number of papers extending these studies for more sample countries or longer sample periods. See e.g. Eijffinger and van Keulen (1995), de Haan et al. (1998), Dvorsky (2000), Sturm and de Haan (2001) and Cukierman et al. (2002).

[11]For a critique of the basic design of these studies see Posen (1993, 1995).

[12]Updates of the turnover rate index can be found in Cukierman et al. (1992), de Haan and Kooi (2000) and Sturm and de Haan (2001).

While the developed indices of legal or actual central bank independence are principally useful, the empirical results based on these indices are nevertheless somewhat questionable. As it was shown in the previous section, both central bank independence and conservatism have an influence on the macroeconomic performance of a monetary authority. Thus, focussing on statutory central bank independence will result in misleading evidence whenever the degree of conservatism differs between countries.[13]

One of the few attempts to measure the degree of conservatism explicitly has recently been made by Berger and Woitek (2005). Their basic approach is to classify the members of the German Bundesbank Council according to their conservatism. In order to do so Berger and Woitek study who nominated the single Council members. The individual members are either selected by the central government or by the local German States. Assuming that both the central or local governments select members sharing their own degree of conservatism and right-wing governments to be more conservative than left-wing ones the authors calculate the average degree of conservatism of the Bundesbank Council over time. Since Berger and Woitek focus exclusively on the German Bundesbank and assume that the degree of central bank independence did not change over the sample period, they end up with a pure measure of conservatism. While this measure of conservatism is properly derived it is less suitable for international comparisons.

A conceptual appealing approach of measuring actual central bank independence was proposed by Eijffinger et al. (1996). The authors argue that the factual degree of central bank independence comes forward in differing structural pressures to lower or raise money market rates. Basically they argue that more independent central banks have lower incentives to stimulate the economy as more

[13] Sometimes, the developed indices of legal central bank independence also include some aspects of conservatism. As an example, the index of legal central bank independence by Cukierman (1992) also refers to the importance of the goal of price stability.

dependent central banks, given the same macroeconomic situation. To uncover these structural differences the authors estimate prime rate reaction functions of 10 central banks within a fixed-effects panel approach thereby using inflation, economic growth and the current account surplus as control variables. The authors then interpret the fixed effects as a measure of average actual central bank independence and find this measure to coincide well with several legal indices of central bank independence.

4 Empirical analysis

4.1 Estimation approach

Our empirical approach of measuring the joint degree of central bank independence and conservatism builds upon the idea of Eijffinger et al. (1996) to study differences in the behavior of a sample of central banks. However, different from Eijffinger et al. (1996) we base our identification method upon a theoretical model. Accoding to the model outlined in section 2 central banks reveal their degrees of effective conservatism by their monetary policies chosen in consequence of occurring unemployment gaps.

In practice monetary authorities appear to alter interest rates in a sequence of small steps to reach the desired interest rate level. In line with most of the literature, we therefore allow for interest rate smoothing in our empirical approach and assume the following dynamic adjustment of interest rates to the optimal level i_t^* (see e.g. Judd and Rudebusch 1998):

$$\Delta i_t = \theta \cdot (i_t^* - i_{t-1}) + \rho \cdot \Delta i_{t-1} = \theta \cdot \Delta^* i_t + \rho \cdot \Delta i_{t-1}. \tag{11}$$

In consequence, the reaction function takes the form:

$$\Delta i_t = -\theta \cdot (\beta^M \cdot (\bar{u} - u^*) + \pi^*) - \theta \cdot \beta^M \cdot \widehat{u}_t + \theta \cdot \pi_{t-1} + \rho \cdot \Delta i_{t-1} + \theta \cdot \eta_t \tag{12}$$

To analyze the degree of effective conservatism we estimate a common interest rate reaction function in a panel framework and allow for differing reaction coefficients for unemployment gaps while - in line with the model's predictions - assuming common coefficients for all additional control variables. Rearranging equation (12) and using n as a country index allows us to separate four determinants of central banks' interest rate decisions. The preferred interest rate variation is the sum of a country-specific effect

$$\lambda_{0,n} = -\theta \cdot (\beta^M \cdot (\bar{u} - u^*) + \pi^*), \tag{13}$$

the lagged effect of the central banks' interest rate policies

$$\lambda_1 \cdot \Delta i_{n,t-1} = \rho \cdot \Delta i_{n,t-1}, \tag{14}$$

the lagged effect of inflation

$$\lambda_2 \cdot \pi_{n,t-1} = \theta \cdot \pi_{n,t-1} \tag{15}$$

and the reaction to the unemployment gap

$$\lambda_3 \cdot \widehat{u}_{n,t} = -\theta \cdot \beta^M \cdot \widehat{u}_{n,t}, \tag{16}$$

with $\beta^M = \beta^G - \bar{\beta} \cdot I$.

Different from Eijffinger et al. (1996) we employ a logit approach to estimate the reaction functions. While principally, interest rates are continuous variables, in practice central banks vary prime interest rates only in discrete steps of 25 basis points. Thus, estimating central banks' reaction functions using OLS would be inappropriate. We therefore follow the practice of Lapp et al. (2003), Gerlach (2004) and Jansen and de Haan (2006) to make use of the methods of categorical data analysis to estimate central banks' reaction functions and employ an ordered logit model for this purpose.[14] In order to do so we first transform the time series

[14] One might also use a probit model for the estimations. However, as it is well-known, the estimation results differ only slightly between the two methods.

of prime interest rate variations into a ternary ordinal variable $\Delta^{\#} i_{n,t}$ which has the value +1 if interest rates were kept constant, +2 if interest rates were raised and 0 if interest rates were lowered. Since prime interest rate variations of more than 25 basis points are relatively rare events this procedure helps to avoid ending up with a scarce number of observations in the respective classes.

Due the discrete steps in the interest rate policy the preferred variation of the prime rate is not directly observable. Thus, we can interpret $\Delta i_{n,t}$ as a continuous latent variable defined by

$$\Delta i_{n,t} = \lambda_{0,n} + \lambda_1 \cdot \Delta i_{n,t-1} + \lambda_2 \cdot \pi_{n,t-1} + \lambda_{3,n} \cdot \widehat{u}_{n,t} + \epsilon_{n,t}. \tag{17}$$

In line with the basic assumption of the logit estimation approach we assume the residual $\epsilon_{n,t}$ to follow a standard logistic distribution. Instead of the preferred variations from equation (17), we observe the discrete variations of the prime rate. The central banks' behavior is consistent with the assumption that the preferred change in the prime rate has to exceed certain threshold levels to be realized. With $\bar{\lambda}_0, \bar{\lambda}_1$ being the unobservable threshold levels (cutpoints) we can express the factual prime interest rate variations as

$$\Delta^{\#} i_{n,t} = \begin{cases} 0 & : \quad \Delta i_{n,t} < \bar{\lambda}_0 \\ 1 & : \quad \Delta i_{n,t} \in (\bar{\lambda}_0, \bar{\lambda}_1) \\ 2 & : \quad \Delta i_{n,t} > \bar{\lambda}_1. \end{cases} \tag{18}$$

The ordered logit approach is based upon the idea to estimate the cumulative probabilities of the different outcomes of the dependent variable. The cumulative probabilities are given by

$$P^{cum}_{j,n,t} = \sum_{k=0}^{j} P(\Delta^{\#} i_{n,t} = k) = P(\Delta^{\#} i_{n,t} \leq j) \tag{19}$$

with j denoting the three classes of possible prime rate variations (lowered, constant, raised). Thus, $P^{cum}_{0,n,t}$ is the probability that the central bank in country n will reduce the prime rate in period t. Analogously, $P^{cum}_{1,n,t}$ is the probability that the prime rate is either reduced or held constant. Based on the cumulative probabilities we can calculate the so-called logits as

$$logit[P^{cum}_{j,n,t}] = ln\frac{P^{cum}_{j,n,t}}{1 - P^{cum}_{j,n,t}}. \tag{20}$$

Using the calculated logits we then estimate the coefficients of the equation

$$logit[P^{cum}_{j,n,t}] = \bar{\lambda}_j - \lambda_{0,n} - \lambda_1 \cdot \Delta i_{n,t-2} - \lambda_2 \cdot \pi_{n,t-1} - \lambda_{n,3} \cdot \hat{u}_{n,t} - \epsilon_{n,t} \tag{21}$$

using the maximum likelihood procedure. The estimated coefficients describe the influence of the right-hand variables on the cumulative logits and thus on the probabilities for the three different prime rate scenarios. By estimating equation (21) we also obtain values for the cutpoints $\bar{\lambda}_0, \bar{\lambda}_1$.

4.2 General estimation issues and data

Our dataset consists of 11 industrialized countries: Belgium, Canada, Denmark, Finland, France, Germany, Spain, Sweden, Switzerland, the United Kingdom and the United States. We make use of monthly data for the period of January 1988 to December 1998. Thus, our sample period covers the pre-EMU era, ending with the takeover of responsibility for monetary policy by the European Central Bank. While the sample period is long enough to have sufficient observations to study central bank behavior it is also short enough to assume reaction coefficients to be stable.

Some of the sample countries took part in the European Exchange Rate Mechanism (ERM). One might suspect that estimating Taylor-type monetary policy rules is misleading in this context because the referring central banks had to chose their prime interest rates to keep their exchange rates under control. However, none of the sample countries ever fixed its exchange rate completely to the anchor currency, the ECU. While the ERM demanded the participating countries to hold their exchange rates within a certain band, the exchange rates were always flexible to a certain extent, thus giving the referring countries at least some room to react on the actual economic situation. While this band was relatively narrow in the beginning (+/- 2,25%) it was widened very soon to +/- 15% which is very close to free floating exchange rates. While the US, Canada, Switzerland and Sweden did not take part in the ERM, Germany was the only sample country which decided to keep its interest rate within the narrow bands of (+/- 2,25%) througout the whole sample period. UK and Italy took part in the narrow bands system from October 1990 until April 1992. Both countries decided to leave the ERM during the ERM crisis of early 1992. Spain, France, Denmark and Belgium switched in 1992 to the wide +/- 15% exchange-rate-bands. Finally, Finland did not take part in the ERM before 1996 and then in the version with wide bands. Summing up, the central banks in our sample were not too much restricted in their monetary policies during the sample period by the ERM mechanism. This conclusion is in line with a number of other papers engaging in estimating Taylor-rules for the EU countries for similar sample periods (see e.g. Gerlach and Schnabel (2000)).

The time series data we use in our empirical study were extracted from the OECD Main Economic Indicators Database. The left-hand variable in our estimation is the change in the prime interest rate. As described earlier we transformed the change in the prime interest rate into a ternary variable with the value of +2 if the prime rate was raised, +1 if the prime rate was kept constant and 0 if the prime rate was lowered.

We estimate the unemployment gap on the basis of seasonally adjusted standardized unemployment rates. Since this time series was not available for Switzerland we use the share of registered unemployed in the labor force in this case.[15]

[15]Since the unemployment rates do not enter the estimations in levels, this procedure seems to be adequate.

While a number of different approaches to estimate unemployment or output gaps were proposed in the literature, there is yet no consensus view on the appropriate method.[16] We employ a Hodrick-Prescott-filter to calculate the unemployment potential (Hodrick and Prescott 1997, Gerdesmeier and Roffia 2004, Adam and Cobham 2004).[17] Furthermore we normalized the gap measures to a mean of 0 and a variance of 1. This seems to be adequate since the variance of the gaps differed considerably between countries. However, the empirical results are quite similar for non-normalized unemployment gaps.[18]

While the simple theoretical framework presented in section 2 is useful to identify the variables in the central bank reaction functions, it can hardly be employed to specify the appropriate lag structure of the interest rate reaction functions. According to the model the optimal reaction is, inter alia, a function of the current unemployment gap. However, due to publication lags of the data on the real economy (e.g. Nelson 2000, p. 13) it might be appropriate to use lagged unemployment gaps. We let the data decide on the optimal lag structure in the interest rate reaction function and employ the Akaike (AIC) and the Bayesian information criterion (BIC) for this purpose. For the unemployment gap we allow lags up to two months. Since our study focuses on the country-specific reaction of central banks to the unemployment gap we allow the lag specification to differ between central banks of the sample countries.

Inflation was calculated on the basis of the consumer price index (all items). According to the model, inflation should enter the interest rate reaction function with a lag. Again, we employed the information criteria to select the optimal lag structure. We allowed for inflations lags of up to 6 months.

[16] See Billmeier (2004) for a review of various approaches.

[17] The gaps were computed from 1979 to 2007 thus avoiding typical start- and endpoint problems of the filter-method. The smoothing parameter in our baseline specification is 100000. However, the results are quite robust with respect to different levels of the smoothing parameter.

[18] The results are available from the authors on request.

All time series of regressors were tested for stationarity. Since we allow for country-specific coefficients of the unemployment gap the evaluation of stationarity of this variable was based on seperate unit-root-tests such as the Augmented Dickey-Fuller-test and the Phillips-Perron-test. For all additional time series we employed various panel-unit-root-tests such as the Levin-Lin-Chu-test, the ADF-Fisher-Chi-Square-test and the Phillips-Perron-Fisher-Chi-Square-test. The results of the unit-root-tests are reported in the appendix. The time series of unemployment gaps turn out to be stationary in all countries and thus can be used in levels in the estimations.[19] Since the hypothesis of the existence of a unit-root can not be rejected for the other variables we use the first difference of these variables in our estimations. The conducted panel unit-root tests indicate that the differenced time-series turn out to be stationary.

4.3 Estimation results

The estimation results are summarized in table 1.[20] Column (1) displays the result for the baseline specification (equation (21)) using the first difference of lagged year-on-year inflation rates and the standardized and normalized unemployment gap. All coefficients have the expected sign. Except for Sweden and Germany the country-specific reaction coefficients of the unemployment gaps are significant at least on a 95%-level. The unconditional probabilities for the three outcomes of the ternary variable can be deduced from the estimated cutpoints. The unconditional probability for a more restrictive monetary policy is 15.2 %, the one for a more expansive monetary policy is 20.2 % and the one for an unchanged monetary policy course is 64.6 %.

In figure 1 we show a country ranking of the average degrees of effective conser-

[19]Using the first difference of the unemployment gaps does not alter the qualitative results.

[20]We do not report the country-specific effects here. The results are available from the authors on request.

Variable	Country	(I)	(II)	(III)	(IV)	(V)	(VI)
$\hat{u}$ (λ_3)	Belgium	-0.40***	-0.38***	-0.43***	-0.45***	-0.38***	-0.47***
		(-3.00)	*(-2.86)*	*(-3.09)*	*(-3.44)*	*(-2.86)*	*(-3.60)*
$\hat{u}$ (λ_3)	Canada	-1.04***	-1.18***	-0.93**	-1.1***	-1.18***	-1.11***
		(-3.10)	*(-3.74)*	*(-2.50)*	*(-3.55)*	*(-3.74)*	*(-3.57)*
$\hat{u}$ (λ_3)	Switzerl.	-0.26***	-0.25***	-0.28***	-0.30***	-0.25***	-0.30***
		((-3.16)	*((-3.14)*	*(-3.33)*	*(-3.69)*	*(-3.14)*	*(-3.67)*
$\hat{u}$ (λ_3)	Germany	-0.19	-0.19	-0.23*	-0.23*	-0.19	-0.22*
		(-1.45)	*(-1.52)*	*(-1.68)*	*(-1.72)*	*(-1.52)*	*(-1.66)*
$\hat{u}$ (λ_3)	Denmark	-0.52***	-0.51***	-0.54***	-0.54***	-0.51***	-0.47***
		(-2.62)	*(-2.68)*	*(-2.67)*	*(-2.91)*	*(-2.68)*	*(-2.59)*
$\hat{u}$ (λ_3)	Spain	-0.72***	-0.69***	-0.74***	-0.76***	-0.69***	-0.75***
		(-3.54)	*(-3.39)*	*(-3.56)*	*(-3.66)*	*(-3.36)*	*(-3.55)*
$\hat{u}$ (λ_3)	Finland	-0.27***	-0.25***	-0.30***	-0.25***	-0.25***	-0.25***
		(-3.09)	*(-2.93)*	*(-3.33)*	*(-2.86)*	*(-2.93)*	*(-2.87)*
$\hat{u}$ (λ_3)	France	-0.68***	-0.69***	-0.69***	-0.77***	-0.69***	-0.74***
		(-3.42)	*(-3.61)*	*(-3.47)*	*(-3.85)*	*(-3.61)*	*(-3.65)*
$\hat{u}$ (λ_3)	Sweden	-0.51	-0.49	-0.50	-0.50	-0.49	-0.50
		(-1.46)	*(-1.41)*	*(-1.40)*	*(-1.43)*	*(-1.41)*	*(-1.44)*
$\hat{u}$ (λ_3)	UK	-0.31***	-0.28**	-0.31***	-0.24**	-0.28**	-0.23**
		(-2.93)	*(-2.52)*	*(-2.78)*	*(-2.13)*	*(-2.52)*	*(-2.09)*
$\hat{u}$ (λ_3)	US	-0.45***	-0.52***	-0.34**	-0.40**	-0.52***	-0.41**
		(-2.67)	*(-3.07)*	*(-2.01)*	*(-2.39)*	*(-3.07)*	*(-2.44)*
Δi_{t-1} (λ_1)	common	0.06	0.08	0.05	0.07	0.08	0.06
		(0.36)	*(0.26)*	*(0.42)*	*(0.37)*	*(0.26)*	*(0.38)*
$\Delta\pi/\Delta E\pi$ (λ_2)	common	6.13	45.91**	300.24	42.11**	45.90**	40.54**
		(0.44)	*(2.40)*	*(0.48)*	*(2.15)*	*(2.40)*	*(2.05)*
Δi_t^w	common	—	—	—	1.67***	—	1.58***
		—	—	—	*(-4.97)*	—	*(-4.97)*
ΔECU	common	—	—	—	—	0.0002	-0.003
		—	—	—	—	*(0.06)*	*(-1.02)*
$D_{92/93}$	common	—	—	—	—		-0.47**
		—	—	—	—		*(-2.34)*
Cutpoint 0 ($\bar{\lambda}_0$)	common	-1.37***	-1.49***	-1.28***	-1.52***	-1.49***	-1.53***
		(-7.81)	*(-8.42)*	*(-7.16)*	*(-8.69)*	*(-8.42)*	*(-8.67)*
Cutpoint 1 ($\bar{\lambda}_1$)	common	1.72***	1.59***	1.92***	1.60***	1.59***	1.61***
		(9.69)	*(9.05)*	*(10.46)*	*(9.23)*	*(9.05)*	*(9.23)*
LR statistic		109.5***	130.9***	103.4***	162.4***	130.9***	167.6***
Pseudo R^2		0.043	0.049	0.044	0.061	0.049	0.063

Significance levels are reported as follows: * for a 90%-significance-level, ** for 95% and *** for more than 99%. We use Huber-White robust estimates. Country-specific effects are not reported. Column (I) is the baseline specification using a HP-filtered normalized unemployment gap with a smoothing parameter of 100000 and lagged year-on-year inflation rates. Columns (II) and (IV) to (VI) use a lead of inflation. In column (III) the results of an estimation using AR(p) forecasts of inflation are displayed. Column (IV) adds international prime rate changes as a control variable to the baseline specification. Column (V) incorporates the exchange rate vis-á-vis the ECU for countries in the ERM. Column (VI) incorporates both, the exchange rate and the international prime rate changes, and adds a dummy-variable for the years 1992 and 1993.

Table 1: Estimation results of the ordered logit model.

vatism. This ranking is based on the change in the probability of a more expansive monetary policy in consequence of an increase in the standardized and normalized unemployment gap of 1 percentage point (pp). According to the estimation results the monetary policy of the Bank of Canada exhibits the lowest degree of effective conservatism over the sample period. An increase in the normalized Canadian un-

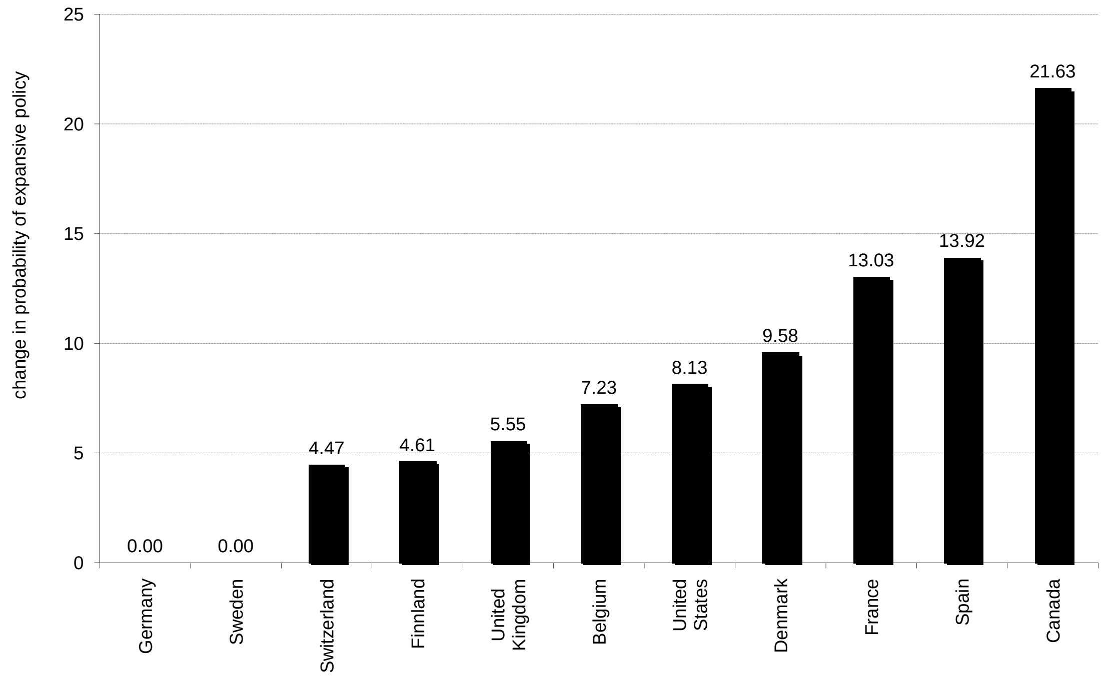

Figure 1: Average effective conservatism of 11 central banks, 1988-1998

employment gap of 1 pp raises the probability of a prime rate cut by 21.63 pp. With 13.92 pp and 13.03 pp we obtain quite large changes in the probability for expansionary policies in Spain and France. We also find comparatively low degrees of effective monetary policy conservatism for Denmark (9.58 pp), the United States (8.13 pp) and Belgium (7.23 pp). Even though there is a significant response to changes in unemployment gaps, the central banks of the United Kingdom (5.55 pp), Finland (4.61 pp) and Switzerland (4.47 pp) show comparatively high degrees of effective conservatism. The estimated coefficients for the central banks of Sweden and Germany turn out to be insignificant (p-value of 0.15 for Germany and 0.14 for Sweden). Hence, neither the German Bundesbank nor Sveriges Riksbank tend to respond to real disturbances throughout the sample period. As a consequence, both central banks have to be considered highly effectively conservative.

In order to find out about the stability of the results, we estimated several variants of the baseline specification. The estimated coefficients for the unemployment gap are quite stable across the various employed specifications of the panel regression. Moreover, the resulting rankings of central banks are almost completely unaffected by the choice of a certain specification. In the following, we briefly describe these specifications.

Since, somewhat surprisingly, lagged inflation turns out to be insignificant in our baseline specification we allowed for a wider range of leads and lags of inflation in specification (II). Due to considerable inside and outside lags in the monetary policy process especially leads of inflation might contribute to enhancing the estimation results. In order to avoid endogeneity problems, we allowed for leads up to a maximum of 12 months. According to previous empirical studies (see e.g. Batini and Nelson 2001) the monetary policy lag is sufficiently long to ensure that this procedure is appropriate. When allowing for leads of inflation, inflation becomes significant with the expected positive sign. However, the estimated coefficients for the unemployment gaps remain almost unchanged.

One might also argue that central banks are foreward-looking with respect to inflation. If so, one might expect current monetary policy to react to medium-term inflation forecasts. However, data on the inflation forecasts of the central banks in our sample were not available. Using factual inflation for appropriate forecast horizons (18 to 24 months ahead) as a proxy for rationally formed medium-term inflation is problematic because of endogeneity problems. We therefore constructed univariate real-time forecasts of inflation. Using a period of 10 years an AR(p) forecast equation for inflation was identified for each point in time t and each central bank n according to

$$\pi_{n,t} = \alpha_{0,n}^t + \sum_{i=1}^{p} \alpha_{i,n}^t \cdot \pi_{n,t-i} + \epsilon_{n,t}. \tag{22}$$

The estimated equation was then utilized to forecast inflation for the next 18 to

24 months.[21] Doing so we obtain forecasts $E\pi_{n,t+h}$ for each country, each forecast horizon h ($h = 18, ..., 24$) and the entire sample period.[22] However, the constructed inflation forecasts (first difference, $\Delta E\pi_{n,t+h}$) turn out to be insignificant in specification (III). While the coefficients slightly change, the ranking of central banks remains stable.

Since specification (II) turns out to be the most reasonable estimation variant as far as inflation is concerned we use this specification as basis for all additional stability tests.

Central banks might be tempted to react to international interest rate trends in order to prevent massive capital outflows and/or to prevent major exchange rate adjustments. We therefore add the change in the interest rate of the G7-countries Δi_t^w to the regression in specification (IV).[23] The coefficient of the G7-interest-rate turns out to be highly significant. However, the ranking of central banks again remains stable in this specification.

Since some of our sample countries took part in the ERM we can not rule out, that interest rate decisions were also influenced by the motive to keep exchange rates within the relevant bands. For ERM countries we therefore added the change in the exchange rate towards the European Currency Unit (ΔECU) as explanatory variable to the regression in specification (V).[24] The coefficient turns out to have the expected positive sign but is not significantly different from zero. The additional

[21] The equations were obtained using Newey-West-Least-Square estimates. Lags were included as long as all autoregressive components remained significant on a 90%-confidence-level.

[22] The appropriate lead structure h in the reaction function was again determined using the information criteria AIC and BIC.

[23] The G7-interest-rate is calculated using a GDP-weighted prime rate of the G7-countries. GDPs are measured in US dollars. The weights were calculated on an annual basis using the end-of-period exchange rate. For the G7 countries in our sample, the national prime rate was excluded when calculating the G7-interest rate.

[24] Due to stationarity issues we again used the first difference of exchange rates. Whenever a country did not take part in the ERM the referring value of the variable is zero, since the existence of the ERM can hardly play a role in this case. The necessary exchange rate data was taken from the IMF database.

results remain virtually unchanged.

Finally, we checked in how far the finding of Gerlach and Schnabel (2000) that interest rates in the EMU countries can be explained by a Taylor-type interest rate rule, except for the period of 1992 to 1993, has an influence on our empirical results. In order to do so, we constructed a dummy variable for this period ($D_{92/93}$) and added it to the regression equation. In fact, the dummy turns out to be significant. However, the unemployment-gap-coefficients and thus the ranking results again remain stable.

5 Summary and conclusions

In this paper we developed an empirical measure of average effective monetary policy conservatism for a sample of 11 industrialized countries over a sample period from 1988 to 1998. Due to the relatively scarce number of observations we did not engage in attempts to relate this measure to the macroeconomic performance of these countries.

Of course, it is interesting to relate our results to those of comparable studies. As we discussed in section 3, a large number of studies was conducted to measure central bank independence and/or conservatism. However, most of these studies are hardly comparable to our results because of the differing sample countries and periods. We therefore concentrate our comparison on the famous index of legal central bank independence of Cukierman (1992).[25]

Cukierman's (1992) index of legal central bank independence covers all countries in our sample. While the original index was developed on the basis of the circumstances in the 1980's and thus earlier than our sample period, a later re-examination (see Cukierman and Lippi 1999) came to the result that the index

[25]It would also be interesting to compare our measure with the results of the panel study by Eijffingeret al. (1996). However, such a comparison is of little value since the two country samples overlap only for 6 countries.

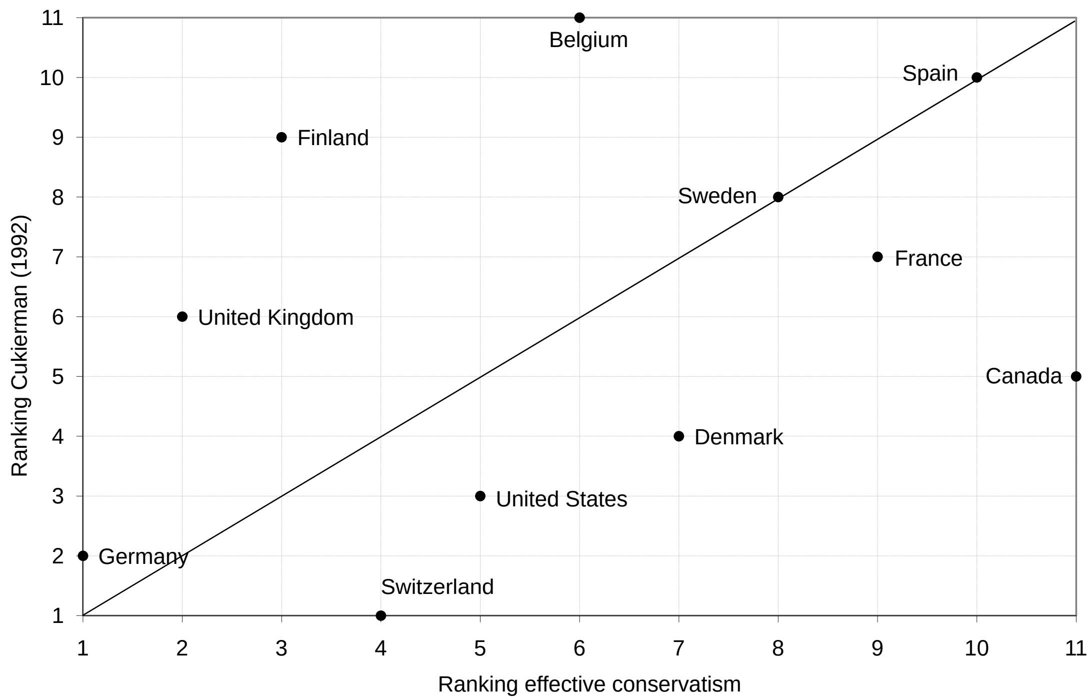

Figure 2: Rank correlation of effective central bank conservatism and legal central bank independence as measured by Cukierman (1992).

remained virtually unchanged in the first half of the 1990s. The rank correlation between our measure and the Cukierman index is shown in figure 2. While Germany, Spain and the United Kingdom are ranked at least very similar under both methods, the ranks of some sample countries differ substantially. Sweden, for example, takes position 8 under the Cukierman methodology but is ranked first in our study. Canada exhibits the least degree of effective conservatism in our study while being ranked fifth by the Cukierman index. Altogether, the rankings reveal that taking into consideration factual independence as well as the median voters' and the central banks' preferences yields a substantially different ranking in comparison to pure legal independence. This heterogeneity indicates that focusing solely on legal central bank independence might lead to severe misinterpretations when trying to explain a country's inflation history.

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A Appendix

Table 2: Results of the single-country unit-root tests.

			test statistic			
Variable	Country	ADF		PP		KPSS
		(I)	(II)	(I)	(II)	(I)
$\widehat{u}_t$ (1)	Belgium	-2.93**	-2.92***	-3.30**	-3.29***	0.051^{+++}
$\widehat{u}_t$ (2)		-4.48***	-4.48***	-4.71***	-4.71***	0.027^{+++}
$\widehat{u}_t$ (1)	Canada	-3.22**	-3.21***	-3.98***	-4.00***	0.045^{+++}
$\widehat{u}_t$ (2)		-4.46***	-4.45***	-5.11***	-5.11***	0.025^{+++}
$\widehat{u}_t$ (1)	Switzerl.	-4.33***	-4.34***	-4.76***	-4.72***	0.060^{+++}
$\widehat{u}_t$ (2)		-5.12***	-5.07***	-3.42**	-3.37***	0.033^{+++}
$\widehat{u}_t$ (1)	Germany	-5.09***	-5.05***	-5.66***	-5.65***	0.042^{+++}
$\widehat{u}_t$ (2)		-6.32***	-6.29***	-6.35***	-6.34***	0.023^{+++}
$\widehat{u}_t$ (1)	Denmark	-3.15**	-3.11***	-3.54***	-3.51***	0.055^{+++}
$\widehat{u}_t$ (2)		-5.96***	-5.94***	-6.34***	-6.34***	0.026^{+++}
$\widehat{u}_t$ (1)	Spain	-2.45	-2.45**	-2.56	-2.56**	0.056^{+++}
$\widehat{u}_t$ (2)		-4.15***	-4.15***	-4.16***	-4.16***	0.034^{+++}
$\widehat{u}_t$ (1)	Finland	-3.18**	-3.17***	-2.95**	-2.91***	0.068^{+++}
$\widehat{u}_t$ (2)		-3.85***	-3.81***	-4.31***	-4.28***	0.054^{+++}
$\widehat{u}_t$ (1)	France	-4.68***	-4.67***	-3.22**	-3.22***	0.046^{+++}
$\widehat{u}_t$ (2)		-6.16***	-6.16***	-4.16***	-4.17***	0.023^{+++}
$\widehat{u}_t$ (1)	Sweden	-2.38	-2.30**	-2.78*	-2.70***	0.082^{+++}
$\widehat{u}_t$ (2)		-3.58***	-3.53***	-5.17***	-5.13***	0.041^{+++}
$\widehat{u}_t$ (1)	UK	-4.99***	-5.00***	-3.79***	-3.77***	0.050^{+++}
$\widehat{u}_t$ (2)		-6.56***	-6.56***	-6.34***	-6.33***	0.027^{+++}
$\widehat{u}_t$ (1)	US	-3.30**	-3.32***	-3.27**	-3.27***	0.045^{+++}
$\widehat{u}_t$ (2)		-4.02***	-4.02***	-4.66***	-4.66***	0.023^{+++}

Abbreviations: ADF–Augmented-Dickey-Fuller, PP–Phillips-Perron and KPSS–Kwiatkowski-Phillips-Schmidt-Shin test. In row (1) the unemployment gap with a smoothing parameter of 100000 is tested. In row (2) the unemployment gap with a smoothing parameter of 14400 is tested. The tests use (I) an exogenous intercept, (II) no exogenous regressors in the test equations. For the ADF and the PP the null hypothesis of a unit root can be rejected * on a 90%-, ** on a 95%- and *** on a 99%-confidence-level. For the KPSS the null hypothesis of stationarity can not be rejected $^{+}$ on a 99%-, $^{++}$ on a 95%- and $^{+++}$ on a 90%-confidence-level.

Table 3: Results of the panel-unit-root tests.

	ADFF		PPF		LLC	
	(I)	(II)	(I)	(II)	(I)	(II)
$\Delta\pi_t$	811.9***	2583.8***	879.2***	2850.9***	-39.08***	-36.54***
$\Delta E_t[\pi_{t+18}]$	801.0***	2450.6***	787.6***	2458.6***	-42.70***	-35.37***
$\Delta E_t[\pi_{t+19}]$	802.0***	2450.5***	788.3***	2458.5***	-43.86***	-35.38***
$\Delta E_t[\pi_{t+20}]$	802.9***	2450.6***	789.5***	2458.6***	-42.94***	-35.41***
$\Delta E_t[\pi_{t+21}]$	803.7***	2450.3***	790.5***	2458.3***	-42.90***	-35.42***
$\Delta E_t[\pi_{t+22}]$	804.4***	2450.4***	791.5***	2458.4***	-42.90***	-35.43***
$\Delta E_t[\pi_{t+23}]$	805.0***	2450.3***	792.4***	2458.3***	-42.88***	-35.45***
$\Delta E_t[\pi_{t+24}]$	805.6***	2450.2***	792.9***	2458.3***	-42.90***	-35.47***
Δi_t^w	136.7***	203.5***	952.9***	2897.3***	-8.44***	-13.25***
ΔECU_t	441.2***	1241.9***	435.6***	1228.3***	-31.54***	-26.39***

Abbreviations: ADFF–Augmented Dickey-Fuller Fisher, PPF–Phillips-Perron Fisher and LLC–Levin, Lin & Chu test. The tests use (I) an exogenous intercept, (II) no exogenous regressors in the test equations. The null hypothesis of a unit root can be rejected * on a 90%-, ** on a 95%- and *** on a 99%-confidence-level. ADFF and PPF assume an individual and LLC a common unit root process.

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