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Measuring Effective Monetary Policy Conservatism

Michael Berlemann
Kai Hielscher

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Autoren / Authors

Michael Berlemann
Helmut Schmidt Universität Hamburg / Helmut Schmidt University Hamburg
Institut für Theoretische Volkswirtschaftslehre / Institute of Economic Theory
Holstenhofweg 85
22403 Hamburg
Germany
Michael.Berlemann@hsu-hh.de

Kai Hielscher
Helmut Schmidt Universität Hamburg / Helmut Schmidt University Hamburg
Institut für Theoretische Volkswirtschaftslehre / Institute of Economic Theory
Holstenhofweg 85
22403 Hamburg
Germany
Kai.Hielscher@hsu-hh.de

Redaktion / Editors

Helmut Schmidt Universität Hamburg / Helmut Schmidt University Hamburg
Fächergruppe Volkswirtschaftslehre / Department of Economics

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Koordinator / Coordinator

Kai Hielscher
wp-vwl@hsu-hh.de

Helmut Schmidt Universität Hamburg / Helmut Schmidt University Hamburg
Fächergruppe Volkswirtschaftslehre / Department of Economics

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MEASURING EFFECTIVE MONETARY POLICY CONSERVATISM

MICHEAL BERLEMANN
KAI HIELSCHER

Zusammenfassung / Abstract

According to the game-theoretic model of monetary policy, inflation is the consequence of time-inconsistent behavior of the monetary authority. The inflation bias can be eased by handing over the responsibility for monetary policy to an independent central bank and appointing a weight-conservative central banker. Countries around the world chose different combinations of central bank independence and conservatism. Most of the existing empirical studies concentrate on measuring legal or factual central bank independence thereby neglecting the degree of conservatism of the monetary authorities. In this paper we show how a joint empirical measure of central bank independence and conservatism can be derived from factual central bank behavior. Based on a panel logit approach we estimate measures of effective monetary policy conservatism for a sample of 11 OECD countries.

JEL-Klassifikation / JEL-Classification: E31, E58

Schlagworte / Keywords: Central Banking, Conservatism, Central Bank Independence, Inflation

1 Introduction

Until the early 1980s monetary policy was typically modeled as a control problem. In the employed models 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. This traditional view eroded in the aftermath of Lucas' [1976] critique on the prevailing macroeconomic policy evaluation methods, trying to predict 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 extend, these doubts evolved as a consequence of the fact that for a long period of time no serious attempts were made to test for 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.

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

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 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.

Maybe the most prominent example for a country relying on the approach proposed by Rogoff [1985] 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 before the academic debate on the importance of central bank independence and conservatism enrolled. 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

[2]See Svensson [1997] or Walsh [2003] for a review of the related literature.
[3]See e.g. Grilli, Masciandaro and Tabellini [1991] or Cukierman [1992].
[4]See e.g. Berger, de Haan and Eijffinger [2001] or Eijffinger, van Rooij and Schaling [1996].

bank independence evolved. Various measures of legal and factual central bank independence were developed and used in empirical studies. 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. 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".

The intention of this paper is to fill the described gap in the literature by studying the average degree of effective monetary policy conservatism of modern central banks. Based on 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 to shocks to the real economy.[5] In order to study the effective degree of conservatism empirically we 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: Section 2 briefly reviews the basic model of time-inconsistency of monetary policy. In section 3 we extend the basic model for the aspects of central bank independence and conservatism and derive a central bank's optimal prime rate reaction function. Section 4 deals with a brief review of the empirical literature on central bank independence and conservatism. In section 5 we present our empirical design, the dataset and the estimation results. The paper ends with a summary of the main findings and some conclusions.

[5]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.

2 The standard game-theoretic model of monetary policy

The stochastic version of the game-theoretic model of monetary policy bases on 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 white-noise random shock. There are two actors in the standard monetary policy game: a government (or a completely government-dependent central bank) controlling inflation via the prime interest rate i_t and trade unions bargaining wages collectively, thereby forming rational expectations on the inflation rate, i.e. $\pi_t^e = E[\pi_t]$. The government is assumed to manipulate its instrument in order to realize both the socially optimal inflation rate π^* and an ambitious employment target $u^* < u^n$. Due to the typical time lags in the transmission process we assume interest rate policy to influence both current and future inflation. Thus, the current inflation rate is determined according to[6]

$$\pi_t = -\gamma \cdot \delta \cdot \Delta i_t - (1 - \gamma) \cdot \delta \cdot \Delta i_{t-1} \tag{2}$$

with $\delta > 0$ and $0 < \gamma < 1$. Some algebraic manipulation allows to reformulate current inflation as

$$\pi_t = \frac{1-\gamma}{\gamma} \cdot \pi_{t-1} - \gamma \cdot \delta \cdot \Delta i_t + \frac{(1-\gamma)^2}{\gamma} \cdot \delta \cdot \Delta i_{t-2}. \tag{3}$$

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

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

[6] For reasons of simplicity we employ a deterministic transmission process. The qualitative results remain unchanged when introducing stochastic disturbances of the transmission process.

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 sequential structure of the game is as follows: First, the government 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 government determines the rate of inflation in order to minimize its expected loss.

The first-best-solution $\pi_t = \pi^*$, $u_t = u^*$ of the described game is infeasible due to the rational expectations of the trade unions. The second-best-solution, i.e. the best solution that can be realized under the trade unions' rational expectations, requires

$$\pi_t = \pi^* + \frac{\beta^G}{1+\beta^G} \cdot \upsilon_t \tag{5}$$

and thus

$$\Delta i_t = \frac{1-\gamma}{\gamma^2 \cdot \delta} \cdot \pi_{t-1} - \frac{1}{\gamma \cdot \delta} \cdot \left(\pi^* + \frac{\beta^G}{1+\beta^G} \cdot \upsilon_t \right) + \frac{(1-\gamma)^2}{\gamma^2} \cdot \Delta i_{t-2}. \tag{6}$$

However, in finitely repeated versions of the time-inconsistency model the government even fails to realize the second-best solution. This is due to the fact that trade unions anticipate the government's incentive to organize surprise inflation as soon as they anticipate the second-best inflation rate. The game ends up in the well-known inflation bias. Formally, the Nash-solution can be calculated by differentiating the government'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

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

Thus, the government will choose

$$\Delta i_t = \frac{1-\gamma}{\gamma^2 \cdot \delta} \cdot \pi_{t-1} + \frac{\frac{(1-\gamma)^2}{\gamma} \cdot \delta \cdot \Delta i_{t-2} - \pi^* - \frac{\beta^G}{1} \cdot (\bar{u} - u^*)}{\gamma \cdot \delta} - \frac{\beta^G}{(1+\beta^G) \cdot \gamma \cdot \delta} \cdot \upsilon_t. \tag{8}$$

Obviously, the inflation bias $\beta^G \cdot (u^n - u^*)$ depends heavily on the relative importance the government (and thus society) assigns to the goals of stable prices and high employment. Ceteris paribus the inflation bias increases in β^G. Note that there is always an inflation bias whenever the government assigns at least a marginally positive relative weight to the ambitious employment target.

Typically, monetary policy is neither a one-shot nor an infinitely repeated game. As Barro and Gordon [1983] showed, reputation formation might contribute to a decrease in the inflation bias in this case. Depending on the reaction of unions to successful cheating by the government the remaining bias is then a weighted average of the socially optimal inflation rate and the inflation bias in the one-shot-game. However, even in infinitely repeated games the inflation bias is not eliminated completely.

3 Central bank independence and conservatism

The Barro-Gordon-model inspired an intensive discussion on how to overcome the time inconsistency problem of monetary policy.[7] 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 and appoint a weight-conservative central banker. In an influential article Rogoff [1985] showed that a conservative central banker who puts less relative weight on the ambitious output target than the government (and the median voter) itself the inflation bias might be eased though not eliminated completely. The loss function of a weight-conservative central banker can be denoted as

$$l_t^C = \frac{1}{2} \cdot (\pi_t - \pi^*)^2 + \frac{1}{2} \cdot \beta^C \cdot (u_t - u^*)^2 \tag{9}$$

with $\beta^C = \beta^G - \bar{\beta}$ and $\bar{\beta} > 0$.

[7] For an overview on the related literature see Svensson [1997].

Since central bank independence is a necessary prerequisite of this solution an extensive literature evolved on the topics how central bank independence could be measured and how these measures are related to inflation and output variability.[8] However, this literature often fails to distinguish properly between the concepts of central bank independence and conservatism.[9]

Following Eijffinger and Hoeberichts [1998] one might describe a monetary authority's loss function by a weighted average of the loss function of the government and the central banker.[10] Defining I as the degree of central bank independence, the monetary authority's loss function can then be denoted as

$$l_t^M = I \cdot l_t^C + (1 - I) \cdot l_t^G. \tag{10}$$

Inserting (4) and (9) in (10) yields the objective function of the monetary authority

$$l_t^M = \frac{1}{2} \cdot (\pi_t - \pi^*)^2 + \frac{1}{2} \cdot (\beta^G - \bar{\beta} \cdot I) \cdot (u_t - u^*)^2, \tag{11}$$

where $\beta^G - \bar{\beta} \cdot I$ 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.[11]

In this setting the monetary authority will choose

$$\begin{aligned} \Delta i_t \;=\; & \frac{1-\gamma}{\gamma^2 \cdot \delta} \cdot \pi_{t-1} \\ & + \frac{\frac{(1-\gamma)^2}{\gamma} \cdot \delta \cdot \Delta i_{t-2} - \pi^* - (\beta^G - \bar{\beta} \cdot I) \cdot (\bar{u} - u^*)}{\gamma \cdot \delta} \\ & - \frac{(\beta^G - \bar{\beta} \cdot I)}{(1 + (\beta^G - \bar{\beta} \cdot I)) \cdot \gamma \cdot \delta} \cdot \upsilon_t. \end{aligned} \tag{12}$$

an end up with the inflation rate

$$\pi_t = \pi^* + (\beta^G - \bar{\beta} \cdot I) \cdot (\bar{u} - u^*) + \frac{\beta^G - \bar{\beta} \cdot I}{1 + \beta^G - \bar{\beta} \cdot I} \cdot \upsilon_t. \tag{13}$$

[8] A survey of this literature can be found in Berger, de Haan and Eijffinger [2001].

[9] For a similar view see Lippi [2000] and Berger and Woitek [2005].

[10] A substantially different interpretation goes back to Lippi [2000]. However, this interpretation is quite different from Rogoff's [1985] original line of argument.

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

Thus, we can conclude that the inflation bias depends inversely on the degree of effective monetary policy conservatism ($\beta^G - \bar{\beta} \cdot I$). 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.[12]

Our model has also implications for interest rate policies of banks with varying degrees of effective conservatism. The stochastic shocks υ_t on unemployment cause fluctuations around the natural level. According to equation (12) central banks will partially offset these distortions via interest rate variations. For a given unemployment shock υ_t the interest rate variation decreases in effective conservatism. Thus, the central banks' reaction to real disturbances reveals information on its effective degree of conservatism.

4 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, Masciandaro and Tabellini [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

[12] This result does not hold in the framework used by Lippi [2000]. However, the major reason for this result is that 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.

legal rules defining the degree of formal central bank independence.[13] 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.[14]

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 Tadeusz [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. Cukierman [1992] proposed two alternative indices of actual central bank independence. The first index is based on the turnover rate of central bank presidents.[15] Basically he argues that a high degree of actual central bank independence should be correlated with a low turnover rate. While this might be a useful measure for developing and transition countries it is somewhat questionable in how far the turnover rate is sufficient to uncover the actual independence of central banks in industrialized countries. That

[13]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, Amtenbrink and Eijffinger [1998], Dvorsky [2000], Sturm and de Haan [2001] and Cukierman, Miller and Neyapti [2002].

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

[15]Updates of the turnover rate index can be found in Cukierman, Webb and Neyapti [1992], de Haan and Kooi [2000] and Sturm and de Haan [2001].

is why Cukierman [1992] developed a second index of factual independence which is based on the results of a large survey study among experts from 24 central banks. The empirical evidence Cukierman [1992] presents points into the direction that the degree of factual central bank independence is negatively correlated with inflation. Furthermore, Cukierman and Webb [1995] construct a political vulnerability index of central banks for the period from 1950 to 1989. The index represents the number political transitions which are followed by the replacement of the central bank governor relative to the total number of transitions. They find significant positive correlations between the indicator of political vulnerability and inflation and inflation variability.

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.[16]

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

[16] 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.

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, van Rooij and Schaling [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 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.

5 Empirical analysis

5.1 Estimation approach

Our empirical approach of measuring the joint degree of central bank independence and conservatism bases on the idea of Eijffinger, van Rooij and Schaling [1996] to study differences in the behavior of a sample of central banks. However, we base our identification method upon the theoretical model outlined in section 3. Given that the structural parameters of the transmission process (γ, δ) are identical in the N different countries, which we will assume in the following, the model predicts that central banks reveal their degree of effective conservatism by their reaction to real disturbances. To analyze this reaction we need a measure of the shock υ_t. 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

$$\upsilon_t = (1 + \beta^G - \bar{\beta} \cdot I) \cdot \underbrace{(u_t - \bar{u})}_{\widehat{u}_t}. \tag{14}$$

Hence, we can substitute υ_t in (12) by equation (14).

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. Different from Eijffinger, van Rooij and Schaling [1996] we employ a logit approach to estimate the reaction functions.

Principally, interest rates are continuous variables. However, in practice central banks vary prime interest rates only in discrete steps of 0.25 base points. Thus, estimating central banks' reaction functions using OLS would be inappropriate. We therefore follow the practice of Lapp, Pearce and Laksanasut [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.[17] In order to do so we first transform the time series of prime interest rate variations into a ternary ordinal variable 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 base points are relatively rare events this procedure helps to avoid ending up with

[17] 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.

a scarce number of observations in the respective classes.

The variation of the prime interest rate preferred by the central bank is given by equation (12). Inserting (14) in equation (12), rearranging 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

$$\lambda_{0,n} = \frac{-\pi_n^* - (\beta^G - \bar{\beta}_n \cdot I_n) \cdot (\bar{u}_n - u_n^*)}{\gamma \cdot \delta}, \tag{15}$$

the lagged effect of the central banks' interest rate policies

$$\lambda_1 \cdot \Delta i_{n,t-2} = \frac{(1-\gamma)^2}{\gamma^2} \cdot \Delta i_{n,t-2}, \tag{16}$$

the lagged effect of inflation

$$\lambda_2 \cdot \pi_{n,t-1} = \frac{1-\gamma}{\gamma^2 \cdot \delta} \cdot \pi_{n,t-1} \tag{17}$$

and the reaction to the unemployment gap

$$\lambda_3 \cdot \widehat{u}_{n,t} = -(\beta^G - \bar{\beta}_n \cdot I_n) \cdot \gamma \cdot \delta \cdot \widehat{u}_{n,t}. \tag{18}$$

In section 3 we deduced this interest rate reaction function for central banks operating in a closed economy. However, in reality central banks might also be influenced by international developments. Since variations of prime rates of foreign central banks cause changes in international interest rate differentials and changing differentials might cause undesired international capital flows central banks might also react to the behavior of other central banks. In order to control for this effect we add the change of the average international prime rate level Δi_t^w as additional possible determinant to the interest rate reaction function.

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

$$\Delta i_{n,t}^* = \lambda_{0,n} + \lambda_1 \cdot \Delta i_{n,t-2} + \lambda_2 \cdot \pi_{n,t-1} + \lambda_{3,n} \cdot \widehat{u}_{n,t} + \lambda_4 \cdot \Delta i_t^w + \epsilon_{n,t}. \tag{19}$$

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 (19), 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

$$\Delta i_{n,t} = \begin{cases} 0 & : \quad \Delta i^*_{n,t} < \bar{\lambda}_0 \\ 1 & : \quad \Delta i^*_{n,t} \in (\bar{\lambda}_0, \bar{\lambda}_1) \\ 2 & : \quad \Delta i^*_{n,t} > \bar{\lambda}_1. \end{cases} \tag{20}$$

The ordered logit approach bases on the idea to estimate the cumulative probabilities of the different outcomes of the dependent variable. The cumulative probabilities are given by

$$P^{cum}_{j,n,t} = \sum_{k=0}^{j} P(\Delta i_{n,t} = k) = P(\Delta i_{n,t} \leq j) \tag{21}$$

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

$$logit[P^{cum}_{j,n,t}] = ln \frac{P^{cum}_{j,n,t}}{1 - P^{cum}_{j,n,t}}. \tag{22}$$

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

$$logit[P^{cum}_{j,n,t}] = \bar{\lambda}_j - \lambda_{0,n} - \lambda_1 \cdot \Delta i_{n,t-2} - \lambda_2 \cdot \pi_{n,t-1} - \lambda_{n,3} \cdot \upsilon_{n,t} - \epsilon_{n,t} \tag{23}$$

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 with which the central banks choose the three different prime rate scenarios. By estimating equation (23) we also obtain values for the cutpoints $\bar{\lambda}_0, \bar{\lambda}_1$.

5.2 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.

All time series of 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 is raised, +1 if the prime rate is kept constant and 0 if the prime rate is lowered. All explanatory variables enter the estimation in a seasonally adjusted version. Whenever a seasonally adjusted time series was not available in the OECD database we employed the ARIMA-X-12 procedure to adjust the data. Monthly inflation rates were calculated on the basis of the seasonally adjusted consumer price index (all items). The change in the world prime rate was calculated using a GDP-weighted prime rate in the G7-countries.[18] For the G7 countries in our sample the national prime rate was excluded when calculating the world prime rate.

We estimate the unemployment gap on the basis of standardized unemployment rates. Since these time series were not available for Switzerland we use the

[18]The GDPs were measured in US dollars. The weights were calculated on an annual basis using the end-of-period exchange rate.

share of registered unemployed in the labor force in this case.[19] While a number of different approaches were proposed in the literature to estimate unemployment or output gaps there is no consensus view on the appropriate method.[20] We employ a Hodrick-Prescott-filter to calculate the unemployment potential (Hodrick and Prescott [1997], Gerdesmeier and Roffia [2004], Adam and Cobham [2004]).[21] Furthermore we normalized the gap measures to a mean of 0 and a variance of 1. This seemed adequate since the variance of the gaps differed considerably between countries. Moreover, this procedure has the advantage that it somewhat controls for a possible violation of the assumption of coinciding structural parameters across countries. However, the empirical results are quite similar for non-normalized unemployment gaps.[22]

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 single unit-root tests such as the Augmented Dickey-Fuller-test and the Phillips-Perron-test. For the time series of inflation and the world prime rate we employed several 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 table 1. According to the stationarity tests the unemployment gaps in all countries are stationary. For all countries the null hypothesis of a unit-root can be rejected at least on a 90 %-confidence level. Since the hypothesis of the existence of a unit-root in the panel of inflation rates could not be rejected we instead used the first difference in

[19]Since the unemployment rates do not enter the estimation in levels this procedure seems to be adequate.

[20]See Billmeier [2004] for a review of various approaches.

[21]The gaps were computed from 1979 to 2007 thus avoiding typical start- and endpoint problems of the filter-method. The smoothing parameter employed here is 100000. However, the estimation results are quite robust with respect to different levels of the smoothing parameter. We tested this for a parameter range from 1000 to 50000, including 14400 as proposed by Hodrick and Prescott [1997].

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

Table 1: Results of the unit-root tests

Variable	Prob.				
	ADF	PP	LLC	ADFF	PPF
Unemployment gap Belgium	0.002***	0.006***			
Unemployment gap Canada	0.008***	0.000***			
Unemployment gap Switzerland	0.000***	0.000***			
Unemployment gap Germany	0.000***	0.000***			
Unemployment gap Denmark	0.004***	0.001***			
Unemployment gap Spain	0.081*	0.078*			
Unemployment gap Finland	0.005***	0.016**			
Unemployment gap France	0.001***	0.004***			
Unemployment gap Sweden	0.033**	0.027**			
Unemployment gap United Kingdom	0.002***	0.000***			
Unemployment gap United States	0.069*	0.0016**			
First difference inflation rate			0.000***	0.000***	0.000***
First difference world interest rate			0.000***	0.000***	0.000***
First difference interest rate			0.000***	0.000***	0.000***

Significance levels are reported as follows: * for a 90%-confidence-level, ** for 95% and *** for more than 99%. Abbreviations: ADF – Augmented Dickey-Fuller-test, PP–Phillips-Perron-test, LLC–Levin-Lin-Chu-test, ADFF–ADF-Fisher-Chi-Square-test, PPF–Phillips-Perron-Fisher-Chi-Square-test. Tests are executed using no exogenous regressors.

the rate of inflation. The panel unit-root tests indicate that the first difference of inflation, of the world prime rate and of the interest rate are stationary.

5.3 Estimation results

While the model presented in section 3 is useful to find out which variables should be included into the estimation of central bank reactions functions, it can hardly be employed to specify the lag structure of central banks' reaction functions. The lag structure depends on factors such as time lags of monetary policy and publication lags of the data on the relevant macroeconomic variables. We therefore let the data decide on the optimal lag structure of the model. The lag or lead structure of the regressors is thus principally based on the results of OLS-regressions of the explanatory variables with different leads or lags on the time series of prime rates. The final decision on the appropriate lead or lag structure is then based on the "Akaike information criterion" (AIC) and "Bayesian information criterion" (BIC). However, there are good reasons to limit the considered leads and lags.

In our model we assumed interest rate changes to affect current and one-period-ahead inflation. Of course, monetary policy affects inflation with a considerably longer time lag in practice (Batini and Nelson [2001]). Thus, the reaction function should be specified in a forward-looking way as far as inflation is concerned.[23] While we can not observe central banks' inflation expectations directly we can use actual inflation as a measure of rationally formed inflation expectations. However, the lead of inflation in the regression must not be too long since otherwise future inflation observations in the reaction function are affected by the current interest rate decision thereby distorting estimation results. When specifying the reaction function we therefore evaluate leads of the inflation variable from 0 to 12 months.

Concerning the unemployment gap we allow here for a backward looking specification of a maximum of 3 months since data on the real economy is typically published with a considerable time lag (Nelson [2000], p. 13). Since our study focuses on the reaction of central banks to real disturbances we allow the lag specification to differ between central banks of the sample countries.

Prime rate changes of central banks are easily and quickly to observe. Thus, if the world prime rate has an effect on the behavior of individual central banks, the effect should be contemporaneous. For past prime rate changes we allow for a lag-specification from 1 to 3 months.

The estimation results are shown in table 2. 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 17.3 %, the one for a more expansive monetary policy is 19.1 % and the one for an unchanged monetary policy course is 63.6 %.

The first 11 rows in table 2 report the country specific reactions of central banks

[23] For similar specifications of monetary policy reaction functions see Clarida, Gali and Gertler [2000], Bernanke and Gertler [1999] or Mehra [1999].

Variable	Country	Coeff.	Std. Err.	z-Stat.	Prob.
Unemployment gap (λ_3)	Belgium	-0.499	0.145	-3.437	0.001***
Unemployment gap (λ_3)	Canada	-1.148	0.330	-3.475	0.001***
Unemployment gap (λ_3)	Switzerland	-0.359	0.097	-3.717	0.000***
Unemployment gap (λ_3)	Germany	-0.192	0.160	-1.204	0.229
Unemployment gap (λ_3)	Denmark	-0.932	0.299	-3.120	0.002***
Unemployment gap (λ_3)	Spain	-0.647	0.171	-3.784	0.000***
Unemployment gap (λ_3)	Finland	-0.213	0.070	-3.061	0.002***
Unemployment gap (λ_3)	France	-0.860	0.214	-4.012	0.000***
Unemployment gap (λ_3)	Sweden	-0.450	0.256	-1.760	0.078*
Unemployment gap (λ_3)	United Kingdom	-0.230	0.096	-2.282	0.017**
Unemployment gap (λ_3)	United States	-0.436	0.174	-2.500	0.012**
Lagged prime rate change (λ_1)	common	0.011	0.051	0.223	0.823
Lead of inflation change (λ_2)	common	43.249	19.240	2.248	0.025**
World prime rate change (λ_4)	common	1.619	0.327	4.956	0.000***
Fixed effect (λ_0)	Belgium	-0.114	0.221	-0.515	0.601
Fixed effect (λ_0)	Canada	-0.324	0.354	-0.914	0.361
Fixed effect (λ_0)	Switzerland	-0.099	0.198	-0.500	0.619
Fixed effect (λ_0)	Germany	-0.087	0.248	-0.349	0.727
Fixed effect (λ_0)	Denmark	-0.105	0.211	-0.499	0.618
Fixed effect (λ_0)	Spain	-0.542	0.267	-2.029	0.042**
Fixed effect (λ_0)	Finland	-0.199	0.199	-1.001	0.317
Fixed effect (λ_0)	France	-0.767	0.262	-2.925	0.003***
Fixed effect (λ_0)	Sweden	-0.525	0.374	-1.404	0.160
Fixed effect (λ_0)	United Kingdom	-0.076	0.227	-0.333	0.739
Cutpoint 0 ($\bar{\lambda}_0$)	common	-1.051	0.165	-6.351	0.000***
Cutpoint 1 ($\bar{\lambda}_1$)	common	1.936	0.168	11.500	0.000***
LR statistic	153.248***				
Pseudo R-squared	0.062				

Significance levels are reported as follows: * for a 90%-significance-level, ** for 95% and *** for more than 99%. We use Huber-White robust estimates of variance.

Table 2: Estimation results ordered logit estimation.

to unemployment gaps. In 10 out of the 11 sample countries the coefficient of the unemployment gap turns out to be significantly negative. This finding implies that the central banks of Belgium, Canada, Switzerland, Denmark, Spain, Finland, France, Sweden, the United Kingdom and the United States will more likely lower prime rates with rising unemployment gaps. Thus, these central banks significantly react to real distortions. Based on the predictions of the model presented in section 3 we interpret this as an indication of a lower degree of average effective central bank conservatism. No significant reaction to unemployment gaps could be detected for Germany. The Bundesbank thus exhibits a high degree of average effective conservatism over the sample period.

The next three lines report the estimation results of the control variables, which are estimated commonly. There is no significant effect of the the lagged variation of the prime rate. However, the change in inflation and the change in the international prime rate both significantly increase the probability that the central banks will raise the prime rate.

The last 10 rows in table 2 report the fixed effects estimations. Only Spain and France significantly differ from the reference country United States.

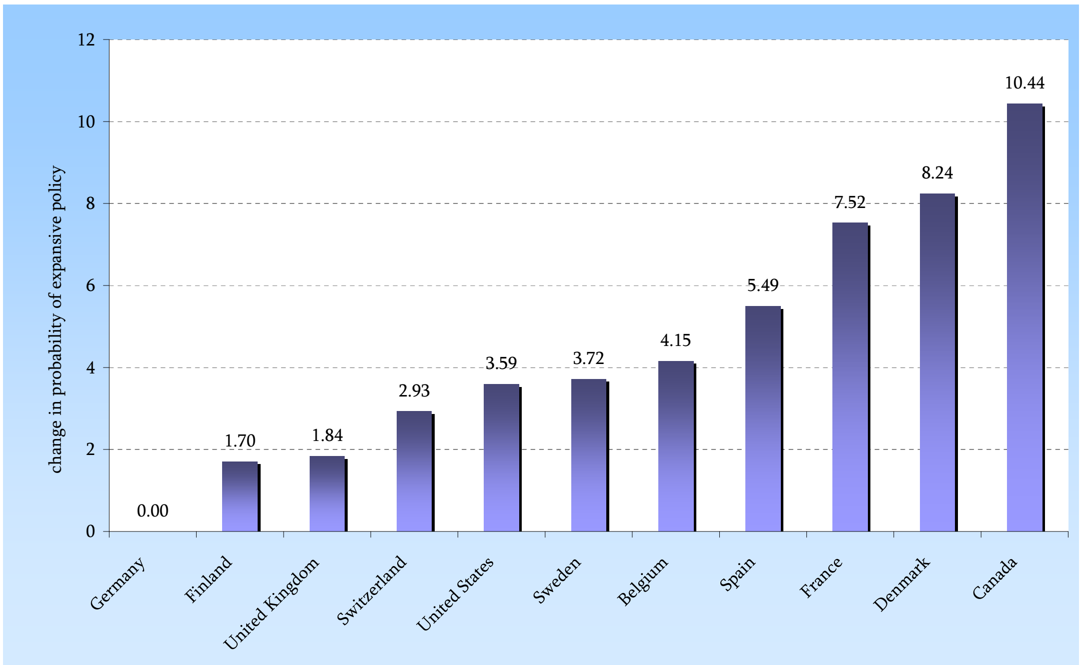

Figure 1: Average effective conservatism of 11 central banks, 1988-1998

In figure 2 we show a country ranking of the average degrees of effective conservatism. This ranking is based on the change in the probability of a more expansive monetary policy in consequence of an increase in the normalized unemployment gap of 50 percentage points (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 unemployment gap of 50 pp raises the probability of a prime rate cut by 10.44 pp. With 8.24 pp and 7.52 pp we obtain quite large changes in the probability for expansionary monetary policies in Denmark and France. We also find comparatively low degrees of effective monetary policy conservatism for Spain (5.49 pp), Belgium (4.15 pp), Sweden (3.72 pp) and the United States (3.59 pp). Even the central banks of Switzerland (2.92 pp), the United Kingdom (1.84 pp) and Finland (1.70 pp) lower prime rates with a higher probability when unemployment gaps occur. However, the three countries show comparatively high degrees of effective conservatism. The German Bundesbank is the only central bank in the sample which does not significantly alter its behavior when unemployment gaps rise.

6 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 4, 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].[24]

[24] It would also be interesting to compare our measure with the results of the panel study by Eijffinger, van Rooij and Schaling [1996]. However, such a comparison is of little value since the

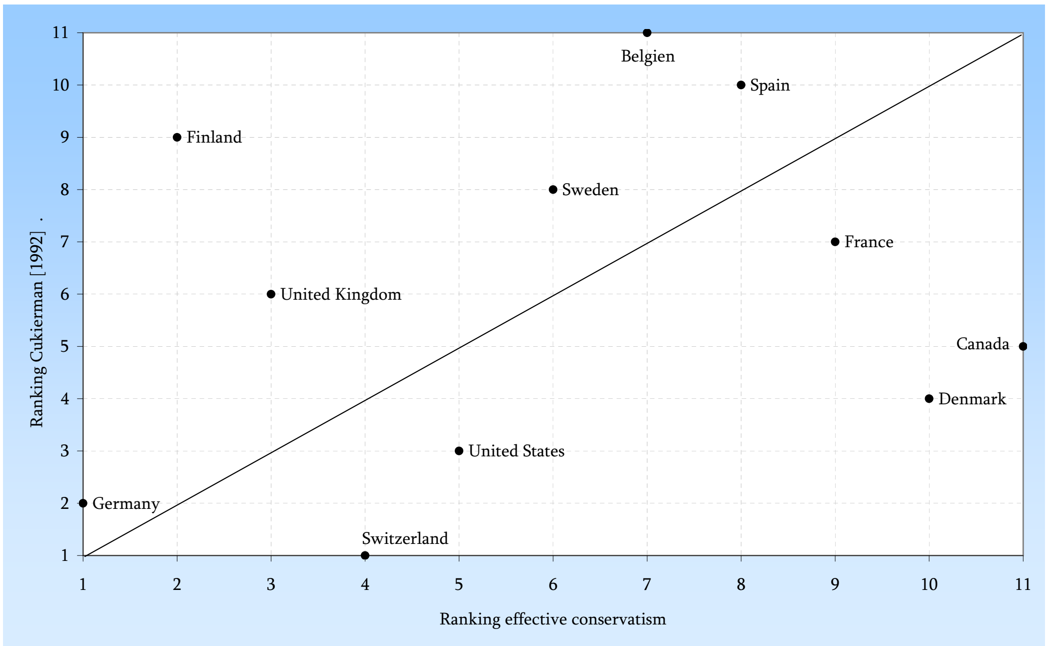

Figure 2: Rank correlation of effective central bank conservatism and legal central bank independence as measured by Cukierman [1992].

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 reexamination (see Cukierman and Lippi [1999]) came to the result that the index 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, France, Sweden and the United States are ranked at least very similar under both methods, the ranks of the other sample countries differ substantially. The most extreme case is Finland ranking second in our study while taking position 9 under 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 le-

two country samples overlap only for 6 countries.

gal 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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