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A TIME-VARYING INDICATOR OF EFFECTIVE MONETARY POLICY CONSERVATISM

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A Time-varying Indicator of Effective Monetary Policy Conservatism

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Zusammenfassung/ Abstract

Based on an extended version of a time-inconsistency model of monetary policy we show that the degree of effective monetary policy conservatism can be uncovered by studying to what extent central banks react to real disturbances. By estimating central bank reaction functions in moving and overlapping intervals for the period of 1985 to 2007 using an ordered logit approach in a panel setting we derive a time-varying indicator of effective monetary policy conservatism for Canada, Sweden, the UK and the US. Employing this indicator we show that increasing effective conservatism tends to lower inflation without increasing the output gap. However, while a higher degree of effective conservatism does not result in lower inflation uncertainty the variance of the output gap tends to decrease.

JEL-Klassifikation / JEL-Classification: E31, E58

Schlagworte / Keywords: central banking, monetary policy, conservatism, central bank independence, inflation

1 Introduction

During the last two decades both economists and politicians devoted considerable interest to the issue of central bank independence. Since the late 1980s many countries around the world decided to increase legal central bank independence. This phenomenon could be observed in well-developed OECD countries such as the United Kingdom or New Zealand, the Eastern-European transition countries but also a number of newly industrializing countries. Two major driving forces of this development played a role herein. On the one hand, countries with highly independent central banks such as Germany or Switzerland showed a good performance in fighting inflation. On the other hand, economic theory delivered important insights in how and why central bank independence might be an important factor in guaranteeing price stability.

The theoretical reasoning behind the importance of central banking bases on work of Kydland and Prescott (1977) and Barro and Gordon (1983). Modeling monetary policy as game between a welfare-maximizing monetary authority and rational wage bargainers Barro and Gordon (1983) show that a government-dependent monetary authority will cause an inflationary bias, i.e. suboptimal high inflation without any real effects. In the aftermath a large literature on how to overcome the problem of time-inconsistent monetary policy evolved.¹ In a seminal paper Rogoff (1985) showed that the inflationary bias can at least be partially offset by appointing a central banker who puts less relative weight on the ambitious output target than the government (and thus society). A necessary precondition for the solution suggested by Rogoff is a perfectly government-independent central bank since a central banker's preferences only matters when there is no possibility to overrule him.

¹See Svensson (1997) or Walsh (2003) for an overview of the related literature.

The outlined theoretical literature soon initiated a large number of empirical studies concerned with the questions, how central bank independence can be measured and whether the various constructed measures of central bank independence are systematically related to the inflation rates in the referring countries. Most of this literature initially focused on legal aspects of central bank independence. Despite the considerable but somewhat inevitable degree of subjectivity in constructing such legal indices² they in general deliver useful measures for an important dimension of central bank independence. However, especially in transition and newly industrializing countries the legal rules were discovered to differ often heavily from the actual practices. Thus, a new strand of the literature started to develop indices of factual central bank independence, for example by calculating the turnover-rates of central bank presidents or conducting expert surveys.³

Altogether, the empirical evidence is somewhat mixed. Typically employing the cross-section regression technique, various studies find a significantly negative relationship between measures of central bank independence and inflation. However, there is also a number of studies contradicting this result.⁴

Various reasons are likely to contribute to the ambiguous empirical results. First, the relevant studies differ considerably with respect to sample countries. Second, the various indices are often calculated for different points in time.⁵ Third, the inevitable subjectivity in constructing the indices and especially in the applied weights of different dimensions of central bank independence often results in quite different measures for the same country. Fourth, almost all empirical studies (at least implicitly) assume that central bank independence is constant over time. Typically, the indicators of central bank independence are calculated for a certain

²See Mangano (1998).

³See e.g. Cukierman (1992).

⁴We review this literature briefly in Section 2.

⁵The same holds true for indices like the Cukierman turnover-index which are constructed over certain periods.

point in time, but are then related to average inflation over long time-spans (of often 10 years or even longer). While this procedure is useful to average out temporary effects in certain countries it is inappropriate when the degree of central bank independence itself is subject to change. In how far this is the case depends on the nature of the indicator. Indices focusing on legal independence only change in consequence of revisions of central bank statutes and thus rather rarely and infrequently. However, factual central bank independence might be subject to more frequent and gradual change. A useful measure of central bank independence should take this into account.

However, the most important reason for the inconclusive results of former empirical studies is the fact that monetary policy outcomes not only depend on the degree of central bank independence but also on the degrees of conservatism of both the median voter and central bank officials.⁶ The reason why this important aspect was neglected in most empirical studies might have to do with the fact that the model employed by Rogoff (1985) does not allow for varying degrees of central bank independence explicitly. A highly useful extension of the model was later delivered by Eijffinger and Hoeberichts (1998). In their extended model the authors show that central bank independence and central bank conservatism are (imperfect) substitutes in reducing the inflationary bias. Moreover, the inflationary bias depends on the median voters' preferences. Consequently, when studying the effect of central bank independence on price stability, it is necessary to control for both the central banks' and the median voters' degrees of conservatism. As an alternative one might also construct a joint measure of a central bank's independence, its conservatism and the median voter's degree of conservatism (effective monetary policy conservatism) and relate it to the inflation performance of the

⁶In some cases the developed indices of legal central bank independence also include aspects of conservatism. As an example, the index of Cukierman (1992) also refers to the importance of the goal of price stability.

referring country.

In particular if central banks are not completely independent from governments, it is likely that the degree of monetary policy conservatism is subject to change in the course of time. According to the median voter theorem a government has to stick to the median voter's preferences in order to ensure to be reelected. As it is shown e.g. in Berlemann (2005), the weights the median voter attaches to the goals of stable prices and high employment might change considerably over time. With an at least somewhat government-dependent central bank this translates into changes of monetary policy conservatism. Moreover, changes in the effective degree of monetary policy conservatism might arise from (infrequent) changes in the central banks' statutes and in the central bank's degree of conservatism (e.g. changes in the governing body of a central bank; see Berger and Woitek, 2005). Thus, when constructing meaningful empirical measures of effective conservatism these measures should be allowed to vary over time and react to changes in central bank independence and the conservatism of both the central bank and the median voter.⁷

In this paper we contribute to the literature by developing a time-varying indicator of effective monetary policy conservatism which is based on the observed behavior of central banks. In order to do so we derive an optimal central bank reaction function from an extended version of the Eijffinger and Hoeberichts (1998) model. We then show that the monetary authority's⁸ degree of effective conservatism can be uncovered by studying to what extent a central bank reacts to real

⁷Note that the fact that effective monetary policy conservatism might change in the course of time also implies that time-invariant measures of central bank independence and/or conservatism (of the central bank or the median voter) are inadequate control variables in panel studies.

⁸In principle, it is a governments privilege to conduct monetary policy. However, most governments around the world decided to delegate the responsibility for monetary policy to more or less independent central banks with differing degrees of conservatism. We therefore use the term "monetary authority" to describe the whole institutional setting determining the conduct of a country's monetary policy.

disturbances. By estimating central bank reaction functions using an ordered logit approach in a panel setting we obtain country-specific reaction coefficients on real disturbances. Based upon these coefficients we define the indicator of effective conservatism as the change in the odds ratios of an expansionary monetary policy which is induced by a real shock. Estimating reaction functions in moving and overlapping intervals for the period of 1985 to 2007 leads to a monthly indicator of effective conservatism of the monetary authorities of Canada, Sweden, the United Kingdom and the United States. Based on this indicator we present empirical evidence in favor of the hypothesis that effective conservatism matters for a country's macroeconomic performance.

The paper is organized as follows: Section 2 briefly summarizes previous empirical evidence. In Section 3 an optimal central bank reaction function is derived from a game-theoretic model of monetary policy. In Section 4 we estimate central bank reaction functions and construct a time-varying indicator of effective conservatism. In Section 5 we use the indicator to evaluate the relation between effective monetary policy conservatism on the one hand and macroeconomic performance on the other. Section 6 brings the paper to its conclusions.

2 A brief review of the existing empirical evidence

The empirical evidence on the conventional view that central bank independence helps to achieve low inflation is somewhat mixed.⁹ Most of this literature is concerned with indices of statutory central bank independence. The majority of these studies finds a significant relation between legal independence and inflation performance (see e.g. Grilli et al., 1991; Cukierman, 1992; Cukierman et al., 1992; Alesina

⁹See Berger et al. (2001) or Hayo and Hefeker (2008) for more detailed reviews of the related empirical literature.

and Summers, 1993; Loungani and Sheets, 1997; Oatley, 1999; Maliszewski, 2000; Cukierman et al., 2002; Gutiérrez, 2003; Jácome and Vázquez, 2005; Carlstrom and Fuerst, 2006; Arnone et al., 2007). However, quite a number of studies present contradicting empirical evidence. As an example, Posen (1993, 1995) does not find central bank independence to enhance inflation performance when controlling for financial sector opposition to inflation. Siklos (2002) supports Posen's argument of a possible endogeneity of independence by showing that central banks that were made more independent in the 1990s already achieved lower inflation in the 1980s. According to Fujiki (1996) the independence-inflation relation is weak when employing panel data and furthermore heavily depends on the sample period. Walsh (1997) finds legal independence to be insignificant in fixed effects models when oil price effects, inflation dynamics and estimates of the natural rate of unemployment are included. Banaian et al. (1998) show on the basis of particular sub-indices of legal independence that the relation between independence and inflation seems to be insignificant or even positive.

However, especially in developing and transition countries factual practices often differ heavily from legal rules (Forder, 1996, 1998). Various studies report anecdotal evidence in favor of this hypothesis (Cukierman, 1992; Hochreiter and Tadeusz, 2000; Berlemann and Nenovsky, 2004). 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) proposes two alternative indices. The first is based on turnover rates of central bank governors.¹⁰ The second indicator is constructed

¹⁰The indicator turns out to be good proxy for actual independence in transition and developing countries while being less suitable for highly developed countries in which the turnover rates differ only slightly and are primarily determined by the legally defined terms of office. For reexaminations of turnover rates see Cukierman et al. (1992), de Haan and Kooi (2000) and Sturm and de Haan (2001).

on the basis of a one-time survey among experts of 24 central banks. Furthermore Cukierman and Webb (1995) construct a political vulnerability index of central banks for the period from 1950 to 1989. It is defined as the number of replacements of the central bank governor in consequence of a political transition relative to the total number of transitions. Altogether, the cited empirical studies find evidence supporting the hypothesis that factual central bank independence improves a central banks' performance in guaranteeing low and stable inflation.

There is also a number of studies focussing on conservatism. Clarida et al. (2000) estimate different forward-looking Taylor-type monetary policy rules for the U.S. Federal Reserve Bank (Fed) and find the reaction coefficient on inflation increased strongly from the pre-Volcker era to the Volcker-Greenspan era. Moreover, this change coincides with an enhanced inflation performance. Favero and Rovelli (2003) also find evidence that shifts in the Feds' preferences contributed to lower inflation and inflation variability.

A conceptual appealing approach of measuring actual central bank independence was proposed by Eijffinger et al. (1996). The authors argue that the actual 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. Their empirical indicator is negatively related to inflation and inflation variability.

While the cited studies are of rather static nature, some measures of central bank independence and conservatism were at least constructed for different time-periods or got updated, later. For example, the Cukierman (1992) index of legal central bank independence has originally been available for 4 different subperiods. Polillo and Guillén (2005) updated the Cukierman index for a large number of central banks for the period of 1990 to 2000 and also considered changes of the legal criteria within this period. Crowe and Meade (2007) replicated the Cukierman index for 2003 using the central bank laws database of the IMF. However, due to the fact that the index focusses on legal independence it changes rather rarely and does neither capture factual independence nor conservatism .

One of the rare studies constructing a time-varying indicator is the one by Berger and Woitek (2005). Focussing on Germany the authors construct a time-varying measure of pure central bank conservatism. In order to do so they analyze the composition of the Bundesbank Council by classifying its single members according to their degree of conservatism. The classification is based on whether the Council member is appointed by a (central or local) right- or left-wing-government assuming right-wing-governments to choose more conservative members than left-wing ones. Assuming the degree of legal independence of the German Bundesbank to remain unchanged over the entire sample period, the authors study the relation between their measure of central bank conservatism and the pattern of inflation. The results support the hypothesis that more conservative Bundesbank Councils generate lower inflation rates and less inflation variability. While the measure of central bank conservatism used by Berger and Woitek (2005) is properly derived it is obviously less suitable for international comparisons.

3 Theoretical framework

3.1 Basic model

Our approach to construct a joint indicator of central bank independence and conservatism bases on the idea by Eijffinger et al. (1996) to estimate monetary authorities' prime rate reaction functions and to derive a measure from the estimated parameters. However, we show that the fixed effect is only a distorted measure of effective conservatism and that it is superior to derive such a measure from a central bank's reaction to real disturbances. In order to do so, we build up on the time-inconsistency literature. Our model stands in the tradition of Barro and Gordon (1983) and Rogoff (1985).¹¹ However, we use the modification proposed by Eijffinger and Hoeberichts (1998) to introduce the degree of central bank independence explicitly into the model.

In the monetary policy game there are two actors: a monetary authority controlling inflation and trade unions bargaining wages collectively, thereby forming rational inflation expectations. The model is based on an expectations-augmented Phillips-curve

$$y_t = \bar{y} + \pi_t - \pi_t^e + \mu_t, \quad (1)$$

where y_t denotes (the log of) output in period t , \bar{y} the (log of) natural rate of output, π_t inflation, π_t^e the wage bargainers' inflation expectations and μ_t a white-noise supply 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

¹¹As Berger and Woitek (2005) state correctly, the implications of effective conservatism regarding 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.

according to

$$\Delta\pi_t = -\Delta i_t + \eta_t \quad (2)$$

with η being a white-noise control error.

Following Eijffinger and Hoeberichts (1998) the loss of a monetary authority (l_t^M) can be expressed as an independence-weighted average of the loss of a country's government (l_t^G) and the loss of the appointed (conservative) central banker (l_t^C)

$$l_t^M = I \cdot l_t^C + (1 - I) \cdot l_t^G, \quad (3)$$

with $0 \leq I \leq 1$. In the case of a completely government-dependent central bank ($I = 0$) monetary policy bases solely on the government's preferences. Under a completely independent central bank ($I = 1$) only the central banker's loss matters.

Both, the government and the central banker are assumed to pursue the goals of stable prices as well as high and stable output with the target levels being π^* and $y^* > \bar{y}$. The loss functions differ solely in the weights assigned to each of the goals. Thus, the one-period loss functions can be denoted as:

$$l_t^i = \frac{1}{2} \cdot (\pi_t - \pi^*)^2 + \frac{1}{2} \cdot \beta^i \cdot (y_t - y^*)^2, \quad (i = C, G) \quad (4)$$

where the relative weight attached to deviations of output from the target level β^G (β^C) denotes the degree of conservatism of the government (the central bank). Following Barro and Gordon (1983) the loss function of the government is assumed to coincide with the social loss function. One might justify this on the basis of the median voter theorem. In the tradition of Rogoff (1985) we assume $\bar{\beta} = \beta^G - \beta^C > 0$ (where $\bar{\beta} > 0$) to denote the difference in the degrees of conservatism of the government and the central bank.

Inserting (4) into (3) leads to the objective function of the monetary authority

$$l_t^M = \frac{1}{2} \cdot (\pi_t - \pi^*)^2 + \frac{1}{2} \cdot \underbrace{(\beta^G - \bar{\beta} \cdot I)}_{\beta^M} \cdot (y_t - y^*)^2, \quad (5)$$

where the relative weight $\beta^M = \beta^G - \bar{\beta} \cdot I$ can be interpreted as the degree of effective monetary policy conservatism.¹² As mentioned earlier it is determined by three (imperfect) substitutes: central bank independence I , the central bank's conservatism β^C and the government's conservatism β^G .

The sequential structure of the monetary policy game is as follows: First, the monetary authority announces an inflation rate. Second, wage bargainers form rational inflation expectations. Third the monetary authority determines the rate of inflation and sets its policy instrument i in order to minimize its loss.

3.2 Nash-Equilibrium

The Nash-solution of the described monetary policy game can be derived by backward induction. Substituting (1) into (5), differentiating with respect to inflation, applying the expectations operator to the first-order condition and solving for expected inflation results in

$$\pi_t^e = \pi^* + (\beta^G - \bar{\beta} \cdot I) \cdot (y^* - \bar{y}). \quad (6)$$

Substituting expected inflation into the first-order condition then leads to

$$\pi_t = \pi^* + (\beta^G - \bar{\beta} \cdot I) \cdot (y^* - \bar{y}) - \frac{\beta^G - \bar{\beta} \cdot I}{1 + \beta^G - \bar{\beta} \cdot I} \cdot \mu_t. \quad (7)$$

Obviously, the inflationary bias $(\beta^G - \bar{\beta} \cdot I) \cdot (y^* - \bar{y})$ as well as inflation variance $\left(\frac{\beta^G - \bar{\beta} \cdot I}{1 + \beta^G - \bar{\beta} \cdot I}\right)^2 \cdot \sigma_\mu$ decrease in the degree of a monetary authority's degree of effective conservatism (i.e. it increases in $\beta^G - \bar{\beta} \cdot I$).

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):

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

¹²Note, that lower values of $\beta^G - \bar{\beta} \cdot I$ correspond to a higher degree of effective conservatism.

The optimal prime rate change of a monetary authority in period t thus depends on (i) a constant term, (ii) the contemporary supply shock and (iii) past inflation. The stochastic shocks μ_t on output cause fluctuations around the natural level. According to equation (8) central banks will partially offset supply shocks via interest rate variations. For a given shock v_t the interest rate variation decreases in effective conservatism.

Output is calculated by inserting equilibrium inflation and inflation expectations in (1):

$$y_t = \bar{y} + \frac{1}{1 + \beta^G - \bar{\beta} \cdot I} \cdot \mu_t. \quad (9)$$

Thus, while an increase in effective conservatism has no effect on the output level, it tends to increase output volatility.

In equilibrium the supply shock can be expressed in terms of fluctuations of output around its natural level, i.e. the output gap \hat{y}_t . Rearranging (9) yields

$$\mu_t = -(1 + \beta^G - \bar{\beta} \cdot I) \cdot \underbrace{(y_t - \bar{y})}_{\hat{y}_t}. \quad (10)$$

Hence, we can substitute μ_t in (8) by equation (10):

$$\Delta^* i_t = -\beta^M \cdot (y^* - \bar{y}) - \pi^* + \beta^M \cdot \hat{y}_t + \pi_{t-1} + \eta_t. \quad (11)$$

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

4 Empirical analysis

4.1 Some general estimation issues

In practice, monetary authorities appear to alter interest rates in a sequence of small steps to reach the desired level. We therefore allow for interest rate smoothing and

assume the following dynamic adjustment of interest rates (actual interest rate change Δi_t) to the optimal level i_t^* (see e.g. Judd and Rudebusch, 1998):

$$\Delta i_t = \theta \cdot (i_t^* - i_{t-1}) + \rho \cdot \Delta i_{t-1} = \theta \cdot \Delta^* i_t + \rho \cdot \Delta i_{t-1}. \quad (12)$$

Furthermore, we take into account the publication lags, especially of variables measuring the real economy. Reliable data is often only available with a delay of several months. Thus, besides the lagged inflation term, we also allow for a possible 'backward-looking' structure of the reaction function with respect to the output gap.¹³

Taking into account the above considerations we can reformulate equation (11) as:

$$\Delta i_t = -\theta \cdot (\beta^M \cdot (y^* - \bar{y}) + \pi^*) + \rho \cdot \Delta i_{t-1} + \theta \cdot \pi_{t-a} + \theta \cdot \beta^M \cdot \hat{y}_{t-b} + \theta \cdot \eta_t. \quad (13)$$

On the basis of (13) and using n as a country index we can separate four determinants of central banks' interest rate decisions. The preferred interest rate variation is the sum of a country-specific effect

$$\alpha_{0,n} = -\theta \cdot (\beta_n^M \cdot (y^* - \bar{y}) + \pi^*), \quad (14)$$

the lagged effect of a central bank's interest rate policy

$$\alpha_1 \cdot \Delta i_{n,t-1} = \rho \cdot \Delta i_{n,t-1}, \quad (15)$$

the lagged effect of inflation

$$\alpha_2 \cdot \pi_{n,t-a} = \theta \cdot \pi_{n,t-a} \quad (16)$$

and the country-specific reaction to the output gap

$$\alpha_{3,n} \cdot \hat{y}_{n,t} = \theta \cdot \beta_n^M \cdot \hat{y}_{n,t-b_n}, \quad (17)$$

with $\beta_n^M = \beta_n^G - \bar{\beta}_n \cdot I_n$ being the country-specific degree of effective conservatism.¹⁴

¹³See section 4.4 for detailed information of the determination of the optimal lead- and lag-structure.

¹⁴Due to the country-specific reaction to output gaps we also allow country-specific lag-structures b_n .

4.2 Basic estimation approach

To analyze the degree of effective conservatism we estimate common interest rate reaction functions in a panel framework. On the basis of equation (13) we thereby allow for differing reaction coefficients for output gaps while - in line with the model's predictions - assuming common coefficients for all additional control variables. Thus, the individual central bank reaction functions differ only in the country-specific effect $\alpha_{0,n}$ (fixed effect) and the reaction to the output gap captured by $\alpha_{3,n}$.¹⁵ The reaction function to be estimated is then given by

$$\Delta i_{n,t} = \alpha_{0,n} + \alpha_1 \cdot \Delta i_{n,t-1} + \alpha_2 \cdot \pi_{n,t-a} + \alpha_{3,n} \cdot \widehat{y}_{n,t-b_n} + \epsilon_{n,t}. \quad (18)$$

In order to estimate (18) one might employ standard panel estimation techniques. However, since in reality prime rates are altered discretely in multiple steps of 25 basis points these techniques are inapplicable here. The preferred prime rate change $\Delta i_{n,t}$ can be interpreted as a continuous latent random variable. Building upon the work of Jansen and de Haan (2005, 2006) and Gerlach (2004) we employ an ordered logit model to analyze central bank behavior assuming the residuals in (18) to follow a standard logistic distribution. The actual prime rate movement $\Delta^{\#}i_{n,t}$ can be defined as a ternary variable

$$\Delta^{\#}i_{n,t} = \begin{cases} 0 & : \Delta i_{n,t} < \lambda_0 \\ 1 & : \Delta i_{n,t} \in (\lambda_0, \lambda_1) \\ 2 & : \Delta i_{n,t} > \lambda_1, \end{cases} \quad (19)$$

with λ_0 and λ_1 denoting the unobservable threshold levels (cut points) for interest rate decisions. Whenever the prime rate is raised above the level of the previous period (i.e., the preferred prime rate change exceeds the threshold level λ_1) the ternary variable $\Delta^{\#}i_{n,t}$ takes the value of 2. When rates are lowered (kept constant)

¹⁵A differing smoothing parameter θ across countries might to some extent limit the cross-section comparability of our measure of effective conservatism. However, since we are primarily interested in the change of the indicator over time it seems to be justified to neglect this potential problem.

it takes the value of 0 (1).¹⁶ Employing the ordered logit technique we can then estimate the *cumulative logits* using the maximum likelihood procedure as

$$\text{logit}[P_{j,n,t}^{cum}] = \lambda_j - \alpha_{0,n} - \alpha_1 \cdot \Delta i_{n,t-1} - \alpha_2 \cdot \pi_{n,t-a} - \alpha_{3,n} \cdot \hat{y}_{n,t-b_n} - \epsilon_{n,t} \quad (20)$$

with $P_{j,n,t}^{cum} = \sum_{k=0}^j P(\Delta i_{n,t} = k) = P(\Delta i_{n,t} \leq j)$ denoting the cumulated probabilities of each category j of the ternary variable and $\text{logit}[P_{j,n,t}^{cum}] = \ln \frac{P_{j,n,t}^{cum}}{1 - P_{j,n,t}^{cum}} = \ln ODDS_{j,n,t}$ describing the (log of) the cumulated odds ratio (*cumulated logit*). The cut points λ_j define the unconditional probabilities of each category of the ternary variable (restrictive, neutral and expansive policy).

4.3 Construction of the time-varying indicator

By estimating (20) we can directly derive a country-specific indicator of effective conservatism (IEC_n) from the reaction coefficients $\alpha_{3,n}$, measuring the extent to which real disturbances influence the probabilities of interest rate variations. We therefore define the indicator as the relative change in the odds ratio (relative to the unconditional odds ratio) of an expansionary policy which is induced by changes in output gaps:¹⁷

$$IEC_n \equiv \frac{ODDS_{0,n}^{\Delta \hat{y}}}{ODDS_{0,n}^{uncond.}} = \frac{\exp(\hat{\lambda}_0 - \hat{\alpha}_{3,n} \cdot \Delta \hat{y})}{\exp(\hat{\lambda}_0)} = \exp(-\hat{\alpha}_{3,n} \cdot \Delta \hat{y}). \quad (21)$$

We should expect a positive value for $\alpha_{3,n}$ since the probability to pursue an expansionary policy (as well as the odds ratio of an expansionary policy) should decrease in consequence of an increasing output gap. In this case we end up with a value of the indicator in between 0 and 1. Obviously, the indicator value will be the smaller

¹⁶While more categories could be taken into account doing so would leave us with a relatively scarce number of observations in each category. We therefore stick to three categories.

¹⁷One could also calculate the change in the probability for an expansionary policy which would be easier to interpret. However, the change in probability does not only depend on the value of the coefficient but also on the unconditional probabilities/cut points. When estimating the reaction function for different time periods, varying estimators for the cut points distort the comparability of indicator-values over time. We therefore use the relative change of the odds ratio.

the more the central bank reacts to real disturbances. Whenever a central bank's interest rate decisions remain unaffected by the output gap, i.e. $\alpha_{3,n} = 0$, we obtain an indicator value of 1. The monetary authority is judged to be completely effectively conservative in this case.¹⁸

A time-varying indicator can be constructed by estimating equation (20) in a panel-setting in moving and overlapping intervals of 60 months. Under this procedure, each month enters the estimation in 60 overlapping intervals. Defining $IEC_n^{t/t+59}$ as the estimated relative odds ratio on the basis of an estimation period reaching from t to $t + 59$, the indicator value in month t , $IEC_{n,t}$, can be defined as

$$IEC_{n,t} = \frac{1}{60} \sum_{i=1}^{60} IEC_n^{t-60+i/t+i-1}; \quad (22)$$

i.e., the unweighted average of the estimated relative odds ratio in the 60 intervals covering month t . This dynamic estimation procedure allows us generating a time-path of effective conservatism in monthly frequency for each sample country. Note that, due to the nature of the indicator, 5 years of observations are lost at the beginning and the end of the sample period.¹⁹

4.4 Data

All employed data is in monthly frequency and was extracted from the OECD database (*Main Economic Indicators*). Our dataset consists of 4 industrialized OECD-countries: Canada, Sweden, the United Kingdom and the United States. The sample period covers January 1985 to June 2007 which yields indicator-values from December 1989 to July 2002.²⁰ While we initially intended to include more countries into our sample we refrained from doing so because of two major reasons.

¹⁸Using a normalized output gap measure we calculate IEC_n on the basis of a change in the output gap of 1 standard deviation. See Section 4.4 for data descriptions.

¹⁹Further reducing the intervals would leave us with a too low number of observations to estimate the reaction coefficients reliably.

²⁰We decided to restrict the sample period to June 2007 since the evolving global financial crisis likely had an influence on the monetary policy decisions in many countries.

First, for various countries the necessary data was not available for the whole sample period.²¹ Second, we only include OECD countries in our sample in which monetary policy was largely unrestrained by exchange rate policy during the sample period.

The ternary variable [see equation (19)] was constructed on the basis of policy interest rates. Year-on-year inflation rates were calculated on the basis of consumer prices (all items). Output gaps were computed on the basis of seasonally adjusted industrial production (excluding construction). While a number of different approaches to estimate output gaps were proposed in the literature, there is yet no consensus view on the appropriate method.²² We employed a Hodrick-Prescott-filter to calculate potential output (Hodrick and Prescott 1997, Gerdesmeier and Roffia 2004, Adam and Cobham 2004). The gap was then calculated as the percentage deviation of production from its potential. Furthermore we normalized the gap measures to a mean of 0 and a variance of 1.²³ The optimal country-specific lag-structure of the output gap (b_n) and the lag-structure of inflation were determined on the basis of the Akaike and the Bayesian information criteria. We allowed for lags up to 6 months in each case.

All regressor variables were tested for stationarity to avoid spurious correlation. Since we allow for country-specific coefficients of the output gap the evaluation of stationarity of this variable is based on single time-series unit-root-tests such as the Augmented Dickey-Fuller test, the Phillips-Perron test and the Kwiatkowski-

²¹While we could include more countries into the sample when shortening the sample period, this comes at the price that the resulting time series of indicator values are short. We then have little possibilities to test the macroeconomic effects of time-varying degrees of effective conservatism. We therefore stick to the 4 sample countries with comparably long sample periods

²²See Billmeier (2004) for a review of various approaches.

²³The gaps were computed from 1980 to 2010 thus avoiding typical start- and endpoint problems of the filter-method. The smoothing parameter in our baseline specification is 100.000. However, the results are very robust with respect to different levels of the smoothing parameter. The empirical results are also quite similar for non-normalized output gaps. See Section 4.6 for robustness checks.

Phillips-Schmidt-Shin test. For all additional time series we employ 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 various unit-root tests are displayed in the appendix (table 1 and table 2). The time series of output gaps turned out to be stationary in all countries. Year-on-year inflation rates and the (lagged) prime rate change appeared to be stationary in the panel. Thus, we used these variables without any further transformations.

4.5 Estimation results

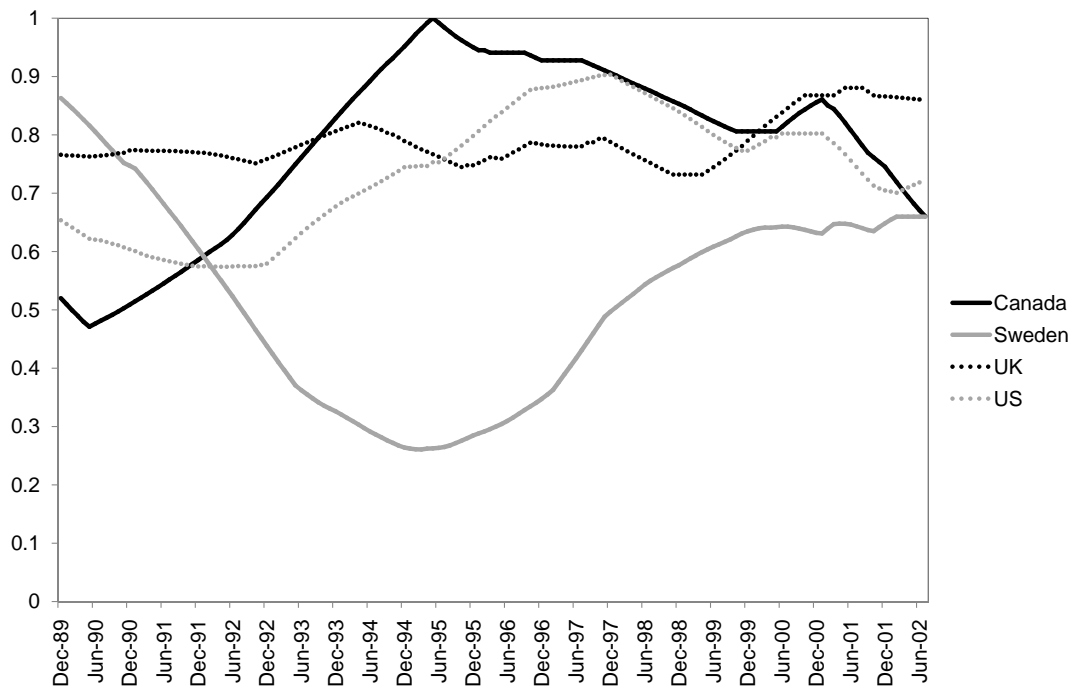
To get a first impression on the overall degree of effective conservatism, we estimated equation (20) over the entire sample period in a first step. The indicator values were calculated on the basis of an increase in the normalized output gap of one standard deviation. In this setting all 4 central banks in our sample reacted significantly to output gaps. However, the Bank of England ($IEC_{UK} = 0.78$) turned out to be the most effectively conservative central bank, followed up by the Federal Reserve ($IEC_{US} = 0.72$). Sveriges Riksbank ($IEC_{Sweden} = 0.64$) and the Bank of Canada ($IEC_{Canada} = 0.63$) turned out to react more actively on real disturbances and thus exhibited lower degrees of effective conservatism.

By applying the methodology described in section 4.3 we then constructed time series of the indicator of effective conservatism $IEC_{n,t}$ for the 4 sample countries. In figure 1 we show the resulting time series of the indicator.

Obviously, the degrees of effective conservatism were subject to considerable change over the sample period. While the indicator values for the United States and the United Kingdom show a comparatively stable development, the values for Canada and especially those for Sweden fluctuate enormously.

In the beginning of the sample period, Sweden shows the highest degree of effective monetary policy conservatism. However, throughout the period of 1989

Figure 1: Development of the indicator of effective monetary policy conservatism, December 1989-July 2002



to 1995 the indicator value decreases from 0.88 to 0.26. In the mid of the 1990s the degree of effective conservatism of Sweden’s monetary policy starts to increase again, thereby reaching a level of 0.65 around 2000. Until the end of the sample period, the indicator remains quite stable.

Quite the opposite development can be observed for the case of Canada. With indicator values of roughly 0.5 Canada exhibits the lowest level of effective monetary policy conservatism in the beginning of the sample period. However, until 1995 the degree to which Bank of Canada reacted to real disturbances decreased steadily. In the mid of 1995 Canada’s monetary policy did not react to output gaps at all, thus reaching an indicator value of 1. However, over the rest of the sample period, the indicator value decreases again, reaching approximately the same level as Sweden in the end of 2002.

The indicator values for the Federal Reserve range between 0.58 and 0.90. While the degree of effective conservatism thus fluctuates quite a bit over the sample period, the development over time turns out to be much more stable than in Canada or Sweden. Although the indicator values show some cycles, they exhibit an upward trend over the sample period.

The indicator values for the United Kingdom remain almost stable for the period of 1989 to 1999. However, throughout the last years of the sample period, a tendency towards a higher degree of effective conservatism can be observed. In the end of the sample period the Bank of England reaches indicator values around 0.85.

4.6 Robustness

To study the robustness of the indicator with respect to the exact specification of the panel regression we conducted several sensitivity tests.

Since central bank reaction functions are often specified in a forward-looking manner as far as inflation is concerned²⁴ we decided to reestimate the model with such a specification. In order to be able to do so, a measure of real-time central banks' inflation expectations is required. In the absence of a suitable dataset we decided to construct time-series of real-time inflation expectations employing a simple forecasting model. Using year-on-year inflation rates (based on the consumer price index) we an AR(p) forecast equation was identified for each point in time t and each central bank n according to²⁵

$$\pi_{n,t} = \lambda_{0,n}^t + \sum_{i=1}^p \lambda_{i,n}^t \cdot \pi_{n,t-i} + \epsilon_{n,t}. \quad (23)$$

²⁴See e.g. Gerlach and Schnabel (2000).

²⁵The length of the sample period to identify each forecast equation was set to 10 years. Thus, the forecast equation e.g. for January 1995 was based on a sample period covering January 1985 to December 1994. The equations were obtained using Newey-West-Least-Square estimates. Lags were included as long as all autoregressive components remained significant on a 90%-level.

The estimated equation was then employed to forecast inflation for the relevant policy horizon of 12 to 24 months. This procedure allowed us to obtain time-series for the real-time inflation expectations for each country, each forecast horizon and the entire sample period. The selection of the appropriate forecast horizon in the reaction function was again based on the Akaike and the Bayesian information criterion.

In a next step, we adapted the model to an open-economy framework. While we derived the estimated reaction function from a closed-economy model, in reality monetary authorities might also be influenced by international developments. For example, variations of prime rates of foreign central banks cause changes in international interest rate differentials which might induce undesired capital flows and exchange rate adjustments. In order to control for the dependence of domestic monetary policy on international interest rate decisions, we added a proxy for the (change in the) international prime rate to the reaction function.²⁶

As explained earlier, the indicator of effective conservatism was derived from the reaction coefficient to the output gap. However, gap measures on the basis of a Hodrick-Prescott-filter are quite sensitive to the applied smoothing parameter. We therefore repeated the estimations using different smoothing parameters ranging from 14.400 as proposed by Hodrick and Prescott (1997) to 1.000.000.

In all these estimation variants the derived indicator values remain virtually unchanged.²⁷ We take this result as an indication that the estimation results are highly robust.

²⁶The international prime rate was calculated as the GDP-weighted prime rate of the G7 countries. GDPs were measured in USD. The weights were calculated on an annual basis using the end-of-period exchange rates. For the G7 countries in the sample the national prime rates were dropped from the aggregation to the international prime rate to avoid deterministic correlation.

²⁷Detailed results are available from the authors upon request.

5 Effective conservatism and inflation performance

The theoretical model outlined in section 3 predicts that increases in effective monetary policy conservatism should decrease both, inflation and inflation variability. Moreover, the model has the implication that increases in effective conservatism should be without any effect on the output level while increasing output fluctuations. Given that we based the derivation of the indicator on the theoretical model we should expect that in fact these hypotheses hold in our 4 sample countries.

In order to study the impact of effective monetary policy conservatism on inflation and output we use the derived country-specific indicators $IEC_{n,t}$ to estimate panel models with fixed country (κ_n) and time effects (v_t). To take into account the relevant outside lag of monetary policy we estimate the panel models for different lag-structures c . In combination with the incorporation of lagged dependent variables ($AR(p)$ -process) this procedure helps to avoid endogeneity problems which might result from the fact that the indicator itself is based on the reaction to the (autocorrelated) macroeconomic variables. The fixed effects panel model has the form

$$Z_{n,t} = \phi_0^c + \phi_1^c \cdot IEC_{n,t-c} + \sum_{i=1}^p \phi_{i+1}^c \cdot Z_{n,t-i} + \kappa_n^c + v_t^c + \epsilon_{n,t}. \quad (24)$$

where Z represents one of the macroeconomic variables: year-on-year inflation (π), inflation variability $Var[\pi]$, the production index (y) and production variability ($Var[\hat{y}]$).²⁸ The coefficient ϕ_1^c denotes the marginal effect of the degree of effective conservatism on the respective macroeconomic variable in c months.

In order to analyze the impact of effective conservatism on inflation and output gap uncertainty we need time-varying proxies of inflation and output gap variance. Following the procedure of constructing the indicator of effective conservatism, we

²⁸The order of the $AR(p)$ process is raised as long as all lags remain significant on a 10 %-level.

calculated variances in intervals of 3 years. The variance measure assigned to month t is then calculated as the unweighted average of the 36 intervals which month t enters. With the sample ending in June 2007, the beginning of the global crisis, this procedure allows us to calculate values of the variance indicators until June 2004.

The results of the estimations of equation (24) for inflation (variability) and output (variability) are shown in figures 2 to 3.²⁹

Figure 2 shows that the marginal effect of effective conservatism on the level of inflation is negative over the relevant policy horizon and significant for lags c ranging from -10 to -18 . Thus, an increase in effective conservatism implies a significant decrease in inflation rates in 10 to 18 months. This result is in line with the models' predictions and the common definition of outside lags of interest rate policy.³⁰ The empirical evidence for inflation variability suggests that the marginal effect of effective conservatism is negative over the relevant policy horizon and significant for long lags of at least 3 years. Hence, higher effective conservatism does not only seem to decrease inflation, but also reduces inflation variability as predicted by the model.

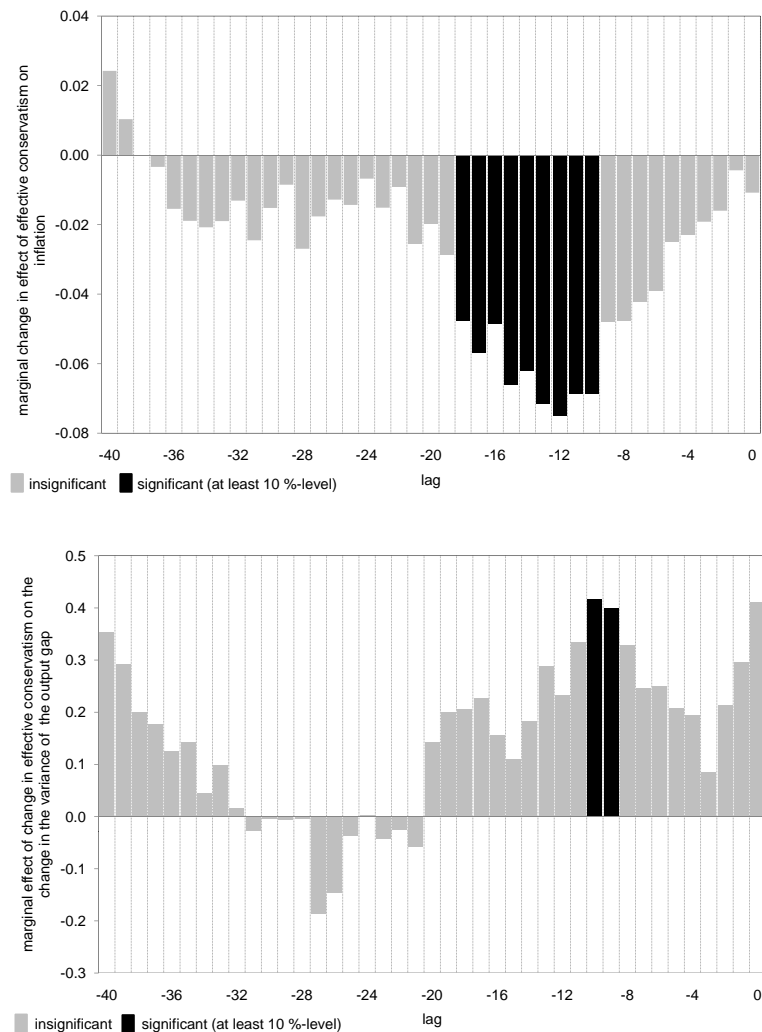
As figure 3 shows, we find no significant effect of effective monetary policy conservatism on production. Again, this is in line with the model's prediction. Finally, we find weak evidence in favor of the hypothesis of a negative impact of effective conservatism on output variability. The displayed results reveal that with comparable short lags of 9 to 10 months, effective conservatism seems to increase production variability.

Summing up, we might conclude that the empirical evidence is in favor of the predictions of the theoretical model we used to construct the indicator of effective

²⁹Since panel-unit-root tests indicate that $IEC_{n,t}$, y and $Var[\hat{y}]$ are non-stationary we used first differences $\Delta IEC_{n,t}$, Δy and $\Delta Var[\hat{y}]$ in these cases. See table 2 in the appendix.

³⁰See Friedman (1961) or Batini and Nelson (2001).

Figure 2: Empirical relation between effective conservatism and inflation (variance): estimated coefficients ϕ_1^c for alternative lags and leads $-40 \leq c$

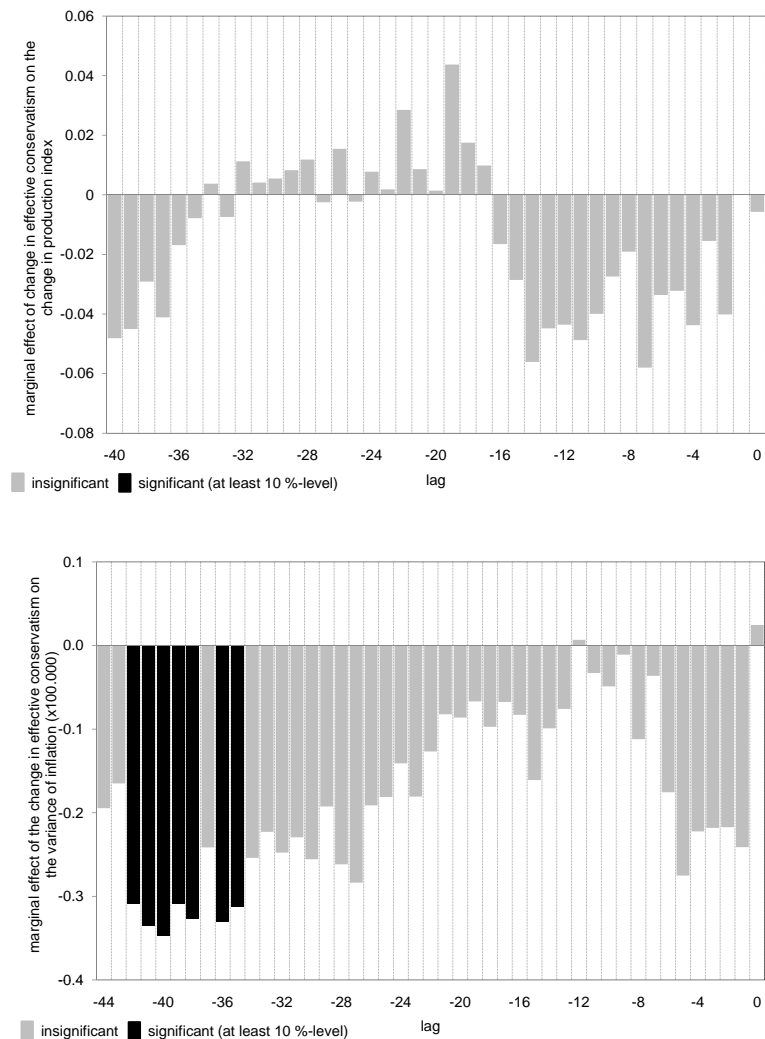


monetary policy conservatism.

6 Conclusions

In this paper we argue that it is a severe shortcoming to attribute inflation (and output) performance of central banks solely to their (formal) degree of independence. Based on a game-theoretic model of monetary policy we showed that besides central bank independence, the degree of conservatism of the median voter as well

Figure 3: Empirical relation between effective conservatism and production (variance): estimated coefficients ϕ_1^c for alternative lags and leads $-40 \leq c$



as the central bank’s degree of conservatism matter for macroeconomic outcomes. We also showed that a central banks’ interest rate reaction function reveals information about the degree of effective monetary policy conservatism. By estimating central bank reaction functions in moving and overlapping intervals in a panel setting we constructed an empirical time-varying indicator of effective conservatism for the monetary authorities of Canada, Sweden, the United Kingdom and the United States from December 1989 to July 2002. The indicator captures the country-

specific extent to which the odds ratios for an expansionary monetary policy are influenced by a real shock in the same country.

The presented empirical results indicate that the degree of effective monetary policy conservatism differs heavily between the sample countries. Much more important: The indicator values fluctuate quite substantially over the sample period. This result is in heavy contrast to the usual assumption in related empirical studies³¹ that the degree of central bank independence (and conservatism) remains stable even over longer sample periods.³²

We also presented empirical evidence for the hypothesis that the contemporary degree of effective conservatism plays a fundamental role for future price stability. Inflation and inflation variability seem to be significantly lower under high degrees of effective monetary policy conservatism while there is no significant effect on the level of production. However, lower inflation and inflation variability is not a free lunch, since higher degrees of effective conservatism seem to increase production variance.

³¹See e.g. Aisen and Veiga (2008) or Adam et al. (2011)

³²This assumption is often an implicit one, e.g. when using one of the various indicators derived in the literature to control for heterogeneity in central bank independence between sample countries.

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A Appendix

Table 1: Results of the single-country unit-root tests.

Variable	test statistic				
	ADF		PP		KPSS
\hat{y} $\lambda = 100,000$	(I)	(II)	(I)	(II)	(I)
$\hat{y}_{Canada,t}$	-4.83***	-4.84***	-4.68***	-4.68***	0.023+++
$\hat{y}_{Sweden,t}$	-6.04***	-6.04***	-12.7***	-12.7***	0.026+++
$\hat{y}_{UK,t}$	-6.50***	-6.51***	-8.61***	-8.62***	0.028+++
$\hat{y}_{US,t}$	-5.24***	-5.25***	-4.47***	-4.48***	0.021+++

For the output gap series we employed single-country unit-root tests as the Augmented-Dickey-Fuller (ADF), the Phillips-Perron (PP) and the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test. 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. Tests were executed for the sample of January 1980 to June 2007.

Table 2: Results of the panel-unit-root tests.

	ADFF		PPF		LLC	
	(I)	(II)	(I)	(II)	(I)	(II)
π_t	48.94***	44.12***	38.69***	44.70***	-4.50***	-5.79***
Δi_t	367.9***	403.2***	386.4***	417.8***	-23.67***	-22.22***
IEC_t	14.06*	3.82	4.70	4.73	-3.31***	-0.08
ΔIEC_t	16.86**	33.15***	17.73**	34.23***	-0.40	-3.92***
$Var[\pi]_t$	13.71*	15.34*	43.04***	47.25***	-0.45	-2.83***
y	3.09	0.64	2.34	0.71	1.23	2.33
Δy	273.7***	316.8***	595.8***	682.3***	-25.01***	-23.24***
$Var[y]$	11.70	9.37	1.31	4.19	-1.58*	-0.61
$\Delta Var[y]$	18.51**	34.22***	28.42***	156.2***	-1.47*	-4.86***

All variables but the output gap were tested for stationarity using panel-unit-root tests as the Augmented Dickey-Fuller Fisher (ADFF), the Phillips-Perron Fisher (PPF) and the Levin, Lin & Chu (LLC) 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. Tests were executed for the sample of January 1980 to June 2007.

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