Cross sectional evidence on the relation between monetary policy, macroeconomic conditions and low-frequency inflation uncertainty

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- Preliminary Draft -

Abstract

In this study, we examine how the interaction between monetary policy and macroeconomic conditions affects inflation uncertainty in the long-term. The unobservable inflation uncertainty is quantified by means of the slowly evolving unconditional variance component of inflation in the framework of the semiparametric Spline-GARCH model (Engle and Rangel, 2008). For a cross section of 13 developed economies, we find that long-term inflation uncertainty is high if central bank governors are perceived as less inflation-averse, if the conduct of monetary policy is rather ad-hoc than rule-based and in economies with a low degree of central bank independence.

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1 Introduction

Inflation uncertainty is costly in terms of aggregate welfare (Fischer and Modigliani, 1976; Barnea et al., 1979). Friedman (1977) and, more recently, Taylor (2012) discuss how certain ways to conduct monetary policy can give rise to IU or macroeconomic uncertainty in general. In a theoretical model, Ball (1992) emphasizes that it is in particular the interaction of monetary policy and inflation which generates IU. Moreover, the importance of the horizon to which uncertainty about future inflation refers is stressed by Ball and Cecchetti (1990).

In this study, we consider a range of policy measures which asses the stance of monetary policy. The relation of these metrics to IU is assessed in terms of a general multi-country model for 13 developed economies which allows for a joint impact of monetary policy and macroeconomic conditions. Although the importance of this interaction for the emergence of IU is theoretically well-established, it has so far been largely disregarded in the related empirical literature. Following the arguments of Friedman (1961) and Ball and Cecchetti (1990), our empirical analysis concentrates on the low-frequency component of IU. This quantity is measured in the framework of the semiparametric Spline-GARCH model proposed by Engle and Rangel (2008). In contrast to conventional GARCH models which are routinely employed to measure IU, this approach enables the specification of a flexible, time-varying low-frequency component of the variance process. We do not consider survey-based measures of IU, which is an alternative approach commonly adopted in the related literature (Zarnowitz and Lambros, 1987; Giordani and Söderlind, 2003). This is because such data is only available for short time periods and a limited set of economies such as the Euro area or the U.S. and therefore precludes the consideration of a larger cross section. This, however, is crucial to identify the impact of changes in the conduct of monetary policy on IU because such changes are usually observed too infrequently within a single economy.

Alternative methods to distinguish between monetary policy schemes are based on the quantification of deviations from the Taylor rule (Taylor 1993), the appointment dates of central bank governors, and an index of central bank independence. In the first case, we obtain a measure of rule-based as opposed to discretionary monetary policy in the sense of Taylor (1993, 2012). Second, the appointment dates serve as a means to separate inflation-averse from other types of central bank governors. We follow Sturm and De Haan (2001) and Dreher et al. (2008, 2010) and relate public perceptions of the inflation-tolerance of monetary policy to the appointment dates of central bank governors. According to Nordhaus (1975), Samuelson (1977), Alesina and Sachs (1988) or Berger and Woitek (2005), liberal governors are typically perceived as more reluctant to adopt disinflation policies. Hence, we regard governors who are appointed under liberal governments as inflation-tolerant as opposed to the presumably inflation-averse governors appointed under other types of governments. Thirdly, a central bank independence index proposed by Sousa (2002) quantifies the degree to which monetary policy in an economy is insulated from political influences.

Our findings are summarized as follows. We document that long-term IU as measured by the Spline-GARCH model is significantly higher during times when monetary policy is more inflation-tolerant than otherwise. Moreover, increases in IU are found if monetary policy is implemented in a rule-based rather than in a discretionary way. In particular, IU increases if governors perceived as less inflation-averse are in power during high-inflation periods. This confirms the theoretical argument of Ball (1992) where high inflation leads to high IU if there is uncertainty about the willingness of central banks to disinflate. This joint effect of inflation and the preferences of monetary policy is markedly stronger than the unconditional link between inflation and IU, which is one of the most widely documented sources of IU in the related empirical literature.

Apart from characterizing monetary policy based on a classification of central bank governors, the (un-)predictability of monetary policy is quantified in terms of deviations from the Taylor rule. We find that IU increases with the degree to which the target interest rate set by monetary policy deviates from the Taylor rule. Notably, IU is higher during periods when the target rate is lower than the prescribed rate, i.e. if monetary policy is overly expansive.

A further influential factor for the emergence of long-term IU is the degree to which governments can influence monetary policy. Relating an index of central bank independence to long-term IU suggests that economies with more independent central banks are, on average, characterized by lower IU.

We analyze these distinct influences jointly in an empirical model which allows for cross-country dependencies and unobserved characteristics of the IU process which vary across economies and time periods. The estimation of a cross-sectional average trend in IU shows that the unconditional IU has been rising since the unfolding of the recent financial- and sovereign debt crisis until the end of the sample period in the year 2010 after it has been decreasing during the *Great moderation* period until the year 2003.

By consideration of distinct methods to approximate IU, we show that the Spline-GARCH-implied measure is most appropriate to examine how macroeconomic determinants and monetary policy are associated with inflation uncertainty. We also document that long-term IU and the interquartile range of surveyed inflation expectations from the Survey of Professional Forecasters of the U.S. FED. are strongly related. As expected, pure measures of inflation variability, which are often associated with IU in the related literature are less suitable as alternative dependent variables since such metrics can only be regarded as noisy approximations of uncertainty. Importantly, by quantifying IU with an ex-post measure such as the the intra-year variability of inflation, we obtain results which indicate a seemingly missing relation between IU, changes in the conduct of monetary policy and the interaction of monetary policy with the level of inflation.

Furthermore, we verify the robustness of our findings with respect to certain subsample choices. First, if IU rises during periods when aggregate uncertainty is high, the high level of overall uncertainty after the begin of the financial- and sovereign debt crisis might lead to spurious findings. Second, the classification of central bank governors could be affected by the membership of six economies from our cross section in the European Monetary Union (EMU), where the European Central Bank (ECB) is in charge of monetary policy. However, excluding observations from the years after 2007 or removing the member states of the EMU from the cross section does not affect our conclusions, which underlines the generality of the documented results.

The remainder of this paper is organized as follows. After a review of the extant empirical literature in Section 2, we introduce our approach to measure IU and describe the empirical setup to examine its potential determinants in Section 3. Next, Section 4 introduces the data set. The empirical results are presented and discussed in Section 5. Finally, Section 6 summarizes the main findings and concludes.

2 Related studies on the determinants of IU

One of the most frequently investigated determinants on IU is the level of inflation. Widely cited discussions of the relationship between inflation and IU are Okun (1971), Friedman (1977) or Fischer and Modigliani (1978). The hypothesis of a causal impact of inflation on IU has been formalized in studies by Devereux (1989) or Ball (1992). In the majority of these theoretical studies, the relation between inflation and IU arises due to the intervention of monetary authorities who respond to changes in either inflation or IU.

Most of the empirical studies test for Granger-causality between inflation and IU. This empirical approach is appealing because of its well-established statistical properties and straightforward interpretation of the corresponding test statistics. However, many of these studies do not explicitly allow for potential effects of distinct monetary policy schemes. Moreover, the joint influence of macroeconomic conditions and the monetary policy framework is typically disregarded.

In several studies, the influence of particular characteristics of monetary policy on IU is examined. In a study on the relation between inflation and IU in the U.S., Evans and Wachtel (1993) point out that that changes in the monetary policy regime might be an important determinant of IU at long horizons. They argue that changes of the monetary policy regime can lead to structural breaks in the inflation process and estimate the timing of regime changes by means of a Markov-switching model based on historical data inflation series. Batchelor and Orr (1991) investigate the effect of inflation targets, the political orientation of the government and other influences on IU in the UK. They employ the root mean squared error (RMSE) computed from a cross section of survey-based expectations of inflation as a measure of IU and find that IU dynamics are driven by changes in the monetary policy environment more strongly than by the level of inflation. However, Batchelor and Orr (1991) point out that their results might depend on particularities of the historical inflation experience of the UK and the properties of the RMSE as a proxy for IU. Similarly, Kontonikas (2004) investigates the relation between IU and inflation targeting in the UK and finds that IU as measured by a GARCH model is lower after the adoption of a formal inflation target by the Bank of England. Caporale and Kontonikas (2009) find that the relation betwen IU and inflation in European economies is affected by the formation of the EMU in the year 1999. Similarly, Hartmann and Herwartz (2013) document that IU is significantly smaller in EMU economies after the Euro introduction as compared to both the situation beforehand and outside the currency union.

Capistrán and Ramos-Francia (2010) or Dovern et al. (2012) study the influence of the monetary policy framework on the cross-sectional dispersion ("disagreement") of survey-based inflation expectations. Disagreement in inflation expectations is often regarded as a measure of IU (Bomberger and Frazer, 1981; Holland, 1993). While Capistrán and Ramos-Francia (2010) find that the disagreement of inflation expectations is lower in economies where inflation targeting strategies are adopted, Dovern et al. (2012) document that the dispersion diminishes with increasing degrees of central bank independence. Similarly, the relation between central bank independence and the level of inflation is studied by Sturm and De Haan (2001) or Posso and Tawadros (2013). Cukierman and Webb (1995) find that the level of inflation typically increases after turnovers before the regular end of governors' term of mandate.

Though the influence of macroeconomic conditions and the characteristics of monetary policy on IU has been documented in several studies, their focus on single determinants of IU such as the level of inflation or the adoption of inflation targeting disregards potentially important joint effects. Such influences are described in central bank reaction functions such as the Taylor rule (Taylor, 1993) or theoretical models as, e.g. Ball (1992), where increasing uncertainty about the ad-hoc or rule-based nature of monetary policy at higher rates of inflation drives IU. In this study, the influence of frequently discussed determinants and the role of their interaction as a source of IU are examined.

3 Measuring and analyzing long-term IU

IU is an unobservable quantity. Which of the many proposed ways to measure it is most suitable, depends on the question under consideration. For example, Ball and Cecchetti (1990) find a more pronounced relation between inflation and low-frequency IU than between inflation and other components of IU. Moreover, the way how monetary policy is conducted can be expected to affect primarily the low frequency movements in IU. Hence, we first consider a model that allows us to separate high- and low-frequency components of IU. In the following, we describe the Spline-GARCH model as our preferred method to proxy IU. Subsequently, we introduce the model which is employed to examine the determinants of long-term IU.

3.1 Empirical strategy

One of the most widely used methods to measure IU is to model the level of inflation in terms of an autoregressive (AR) specification or a reduced-form Phillip curve (Canova, 2007; Stock and Watson, 2008) and to employ the conditional volatility of the corresponding disturbance process, specified in terms of a (G)ARCH model, as an expression of

IU (Engle, 1982, 1983; Bollerslev, 1986).[1] Since in a stationary GARCH model the unconditional variance, i.e. long-term IU, is constant by assumption, we can think of the corresponding conditional variance as a proxy of short-term IU. While this measure may properly reflect the influences of temporary movements in inflation on IU, modeling the response of IU to (permanent) changes in economic policy or macroeconomic conditions (as discussed in Ball and Cecchetti, 1990, or Ball, 1992) requires a specification which allows for secular variations in IU. Hence, in this study we consider the Spline-GARCH model of Engle and Rangel (2008) which separates the conditional variance of inflation into a short-term and a long-term component which smoothly changes over time.

A focus on the slowly evolving component of IU might suggest to use only, say, annual observations for the empirical analysis. However, estimation of the parameters of a Spline-GARCH model based on observations sampled at the yearly frequency is presumably inefficient.

Thus, we adopt a procedure similar to the one employed by Engle and Rangel (2008). We firstly estimate the coefficients of the Spline-GARCH model based on monthly observations and then aggregate the long-term component to a yearly frequency. Next, the implied IU measure is related to indicators of institutional conditions and economic quantities for which only annual observations are available.

3.2 Estimating IU by means of the Spline-GARCH model

We specify the conditional mean of the inflation process as a reduced-form Phillips curve. The inflation rate in economy i, $i = 1, ..., N$, observed in year t, $t = 1, ..., T$, and month m, $m = 1, ..., M$, is denoted by $\pi_{i,t,m}$. Similarly, the growth rate of industrial

[1] In the following, we use the terms *volatility* and *variance* interchangeably.

production is denoted by $y_{i,t,m}$. The country specific reduced-form Phillips curve reads as

$$\pi_{i,t,m} = \nu_i + \sum_{p=1}^{P_i^{(\pi)}} \phi_{i,p}\pi_{i,t,m-p} + \sum_{p=1}^{P_i^{(y)}} \varphi_{i,p} y_{i,t,m-p} + u_{i,t,m}, \tag{1}$$

where ν_i represents a constant, $\phi_{i,p}$ and $\varphi_{i,p}$ are the parameters on lagged inflation and output. The orders of the lag polynomials in $\pi_{i,t,m}$ and $y_{i,t,m}$, respectively, are denoted $P_i^{(\pi)}$ and $P_i^{(y)}$ and are selected by the BIC. The maximum lag order is set to twelve.[2] Furthermore, to keep the notation tractable, we do not account for cases when lag polynomials are covering the instances $m-p, m-p+1, ..., m$ which pertain to distinct years such as $t-1$ and t, say.

Modeling the Phillips curve relation based on output growth instead of unemployment is a commonly adopted way to specify the conditional mean of the inflation process in the empirical literature on IU (Fountas and Karanasos, 2004; Grier et al., 2004). Proceeding in this way seems warranted given the empirically documented stable relation between output growth and unemployment (Blinder, 1997).[3]

We assume that the innovations to inflation are given by

$$u_{i,t,m} = \sqrt{h_{i,t,m}} Z_{i,t,m}, \quad Z_{i,t,m} \overset{iid}{\sim} \mathcal{N}(0,1), \tag{2}$$

$$\text{with } h_{i,t,m} = \tau_{i,t} g_{i,t,m}, \tag{3}$$

where $\tau_{i,t}$ and $g_{i,t,m}$ denote the low- and high-frequency components of the conditional variance. While $g_{i,t,m}$ changes at the monthly frequency and is intended to capture the transitory component of inflation volatility, $\tau_{i,t}$ changes at the yearly frequency only and

[2]Alternative choices for the maximum lag order or the selection of $P_i^{(\pi)}$ and $P_i^{(y)}$ by means of the AIC leads to qualitatively equivalent results.

[3]Moreover, for the economies we examine, monthly unemployment series are not available in the early years of the sample period.

reflects long-term influences such as changes in the institutional conditions of monetary policy.[4] The long-term trend in IU is modeled as an exponential spline function given by

$$\tau_{i,t} = \kappa_i \exp\left(\omega_{i,0}t + \sum_{k=1}^{K_i} \omega_{i,k}(\max(t - t_{k-1}, 0))^2\right). \tag{4}$$

In (4), the flexibility of the trend function increases with the order K_i. Short-term IU is expressed in terms of a unity GARCH process, which reads as

$$g_{i,t,m} = (1 - \alpha_i - \beta_i) + \alpha_i \left(u_{i,t,m}^2/\tau_{i,t-1}\right) + \beta_i g_{i,t,m-1} \tag{5}$$

with $\alpha_i > 0$, $\beta_i \geqslant 0$ and $\alpha_i + \beta_i < 1$. The specification in (5) ensures that $\mathbf{E}[g_{i,t,m}] = 1$. Hence, the time-varying unconditional variance of the innovations to the inflation process is given by $E[u_{i,t,m}^2] = E[g_{i,t,m}\tau_{i,t}Z_{i,t,m}^2] = \tau_{i,t}$, i.e. by the slowly evolving long-term component. Note that the Spline-GARCH model reduces to the standard GARCH model when $\tau_{i,t} = \tau_i$ is constant. The parameters $(\alpha_i, \beta_i, \kappa_i, \omega_{i,0}, ..., \omega_{i,K_i})'$ are estimated by means of quasi-maximum likelihood. The BIC guides the selection of K_i.

We base our measure of IU on the monthly series of conditional variances $h_{i,t,m} = \mathbf{E}_{i,t,m-1}[u_{i,t,m}^2]$, where the expectation is conditional on the information available up to month $m-1$. The $h_{i,t,m}$ can thus be considered as an *ex-ante* measure of the monthly IU. When examining the linkages between IU and macroeconomic and institutional settings, we focus on the dynamics of IU at the annual frequency. For this, we define our measure of annual IU in country i as the square root of the aggregated monthly

[4]Note that our specification is different from the one considered in Engle and Rangel (2008) who assume that both components vary at the same frequency.

conditional variances in year t:

$$IU_{i,t} = \left(\sum_{m \in t} h_{i,t,m} \right)^{1/2} = \tau_{i,t}^{1/2} \left(\sum_{m \in t} g_{i,t,m} \right)^{1/2} . \tag{6}$$

Since $g_{i,t,m}$ is one on average, $IU_{i,t}$ will vary around the slowly moving long-term unconditional volatility of inflation. However, during turbulent times with persistent variations in the short-term component, $IU_{i,t}$ might considerably deviate from $\tau_{i,t}^{1/2}$.

Furthermore, we construct an *ex-post* measure of the intra-annual variability of inflation as

$$SD_{i,t}(\pi) = \left(\sum_{m \in t} (\pi_{i,t,m} - \bar{\pi}_{i,t})^2 \right)^{1/2} , \tag{7}$$

with $\bar{\pi}_{i,t} = (1/12) \sum_{m \in t} \pi_{i,t,m}$. Besides being an ex-post measure of inflation variability, we can think of $SD_{i,t}(\pi)$ as a noisy proxy of $IU_{i,t}$ (see Engle et al., 2013, and Conrad and Loch, 2013). In analogy to (7), we calculate the annual variability of output, denoted by $SD_{i,t}(y)$.

3.3 Determinants of IU

Apart from uncertainty regarding future monetary policy and macroeconomic factors, $IU_{i,t}$ can be driven by various other factors which are mostly outside the range of decision taking at the national level. The importance of such factors is reflected in the debate on the sources of the *Great moderation*. It is highly controversial whether the attenuation of first- and second-order inflation dynamics in many economies during the 1980s and 1990s should be primarily regarded as a success of monetary policy or as the result of a reduced magnitude of inflationary (e.g. oil price-) shocks at a global scale. On the one hand, Taylor (2012) argues that a rule-based type of monetary policy should be ascribed primary responsibility for the *Great moderation*. Empirical support for this argument is provided by Galí and Gambetti (2009), Herrera and Pesavento (2009) or Conrad and Eife (2012). On the other hand, Benati (2008) argues that changes in the type of inflation surprises may be the primary source of the *Great moderation* in the UK. Similarly, Ciccarelli and Mojon (2010) find that a main component of inflation rate fluctuations in the G7 is a common international trend which they refer to as "global inflation".

The presence of global influences on $IU_{i,t}$ might give rise to biases in single-economy time-series estimates regarding the impact of the monetary policy framework on $IU_{i,t}$. Thus, to account for such threats to the validity of the empirical design, we complement the information drawn from the country specific time series by cross sectional data from 13 advanced economies. Following Engle and Rangel (2008), we estimate the relation between $IU_{i,t}$ and its covariates in the framework of the seemingly unrelated regressions (SUR) model. This framework allows us to control for both unobserved heterogeneity and dependencies across economies. The model specification for economy i in year t is given by:

$$IU_{i,t} = x'_{i,t-1}\delta + D'_{i,t-1}\gamma + e_{i,t}, \tag{8}$$

$$\text{where} \quad e_{i,t} = \lambda_t + \eta_i + v_{i,t}, \tag{9}$$

$$v_{i,t} = \rho_i v_{i,t-1} + \epsilon_{i,t}, \tag{10}$$

and $(\epsilon_{1,t}, ..., \epsilon_{N,t})' \overset{iid}{\sim} (0, \Sigma)$. In (8), predetermined macroeconomic quantities are summarized in $x_{i,t-1} = (\pi_{i,t-1}, y_{i,t-1}, SD_{i,t-1}(\pi), SD_{i,t-1}(y))'$. It is theoretically well established and empirically documented that $IU_{i,t}$ increases during periods of higher inflation

(Okun, 1971; Friedman, 1977; Conrad and Karanasos, 2005). Mankiw et al. (2003) discuss the relation between $y_{i,t-1}$ and the disagreement of survey expectations of inflation. Further empirical investigations of this relation are provided by Brunner (1993) or Apergis (2004). In line with these studies, we include past inflation $\pi_{i,t-1}$ and output growth $y_{i,t-1}$ as potential determinants of $IU_{i,t}$. Moreover, the relation between $IU_{i,t}$ and the variability of $\pi_{i,t-1}$ and $y_{i,t-1}$ is examined by including $SD_{i,t-1}(\pi)$ and $SD_{i,t-1}(y)$. A positive relation between the disagreement of inflation expectations and $SD_{i,t-1}(\pi)$ is discussed, e.g., by Capistrán and Timmermann (2009).

Next, we introduce several metrics which quantify the potential effects of predetermined influences on $IU_{i,t}$. These determinants are summarized in the vector $D_{i,t-1}$. The first measure is an index of central bank independence, denoted $indep_i$, which is provided by Sousa (2002). Grier and Perry (1998) or Dovern et al. (2012) study the influence of central bank independence on $IU_{i,t}$ and the disagreement of inflation expectations, respectively. Furthermore, we employ the *Taylor rule* (Taylor, 1993) as a means to quantify the predictability of monetary policy. We specify economy i's target interest rate $R^{\star}_{i,t-1}$ as a function of the real interest rate r_i, the deviation of $\pi_{i,t-1}$ from its target level $\pi^{\star}_i$ and the output gap $\tilde{y}_{i,t-1}$ such that

$$R^{\star}_{i,t-1} = r_i + \gamma_\pi(\pi_{i,t-1} - \pi^{\star}_i) + \gamma_y \tilde{y}_{i,t-1}. \tag{11}$$

In (11), γ_π and γ_y denote the weights attached to deviations of inflation and output from their target- respectively long-run value. Though not all economies in the cross section have explicitly announced inflation targets, inflation targeting as a monetary policy rule has become widespread among industrialized economies during recent decades.

As in Taylor (1993), we set $\gamma_\pi = 1.5$, $\gamma_y = 0.5$ and a level of 2% for both r_i and $\pi_i^\star$. Deviations of the interest rate from the theoretically prescribed target are given by $\tilde{R}_{i,t-1} = R_{i,t-1} - R_{i,t}^\star$. Based on $\tilde{R}_{i,t-1}$, we include the magnitude of past deviations from the Taylor rule, $|\tilde{R}_{i,t-1}|$, in $D_{i,t-1}$. However, $IU_{i,t}$ is not necessarily affected by positive and negative deviations in the same way. Thus, in an alternative model, $D_{i,t-1}$ contains $\tilde{R}^+_{i,t-1} = \tilde{R}_{i,t-1} \times \mathbb{1}(\tilde{R}_{i,t-1} > 0)$ and $\tilde{R}^-_{i,t-1} = \tilde{R}_{i,t-1} \times \mathbb{1}(\tilde{R}_{i,t-1} < 0)$, where $\mathbb{1}(\cdot)$ denotes the indicator function. Thereby we allow for a differential effect of contractionary, respectively expansionary monetary policy. Moreover, the variability of interest rates might reflect a lack of predictability in the conduct of monetary policy. Based on a quarterly series of interest rates $R_{i,t,q}$, we compute the intra-year variability in year $t-1$ as

$$Vr_{i,t-1}(R) = \sqrt{\sum_{q \in t-1} (R_{i,t-1,q} - R_{i,t-1,q-1})^2}. \tag{12}$$

Rudebusch (2002) or Söderlind et al. (2005) discuss the relation between the predictability of monetary policy and distinct forms of inertia in central banks' interest rate setting behavior. We consider metrics such as $\tilde{R}_{i,t-1}$ and $Vr_{i,t-1}(R)$ as a quantitative measures of the monetary policy stance. However, the identification of periods where the most pronounced deviations occur allows to draw a more clear-cut, i.e. a binary distinction between rule-based and ad-hoc ways to conduct monetary policy. For example, Nikolsko-Rzhevskyy et al. (2013) adopt such a strategy by detecting structural breaks in Taylor rule deviations to discern predictable from discretionary monetary policy schemes. An alternative binary classification can be obtained by thinking of monetary policy regimes as rather inflation-tolerant or inflation-averse. We separate these two sorts of conducting monetary policy by means of the dummy variable $d_{i,t-1}^{tolerant}$, which is specified such that $d_{i,t-1}^{tolerant} = 1$ if central bank governors[5] are appointed under left-wing governments $d_{i,t-1}^{tolerant} = 0$ in all other cases[6]. This specification is based on the theoretical model in Ball (1992), where liberal central bank governors give rise to $IU_{i,t}$ because, in contrast to other governors, their determination to disinflate during high-inflation periods is unknown in advance. In this model, $IU_{i,t}$ is triggered by the appointment of less inflation-averse central bank governors only if inflation exceeds a certain threshold level. Thus, in addition to $d_{i,t-1}^{tolerant}$, $D_{i,t-1}$ includes the indicator variable $d_{i,t-1}^{\pi>\mathcal{T}_i}$, where $d_{i,t-1}^{\pi>\mathcal{T}_i} = 1$ if $\pi_{i,t-1}$ is larger than a country-specific threshold value $\mathcal{T}_i$. However, it is unlikely that a particular threshold level is suitable for all economies in the sample. To account for idiosyncratic thresholds across economies, we adopt a data-driven selection of $\mathcal{T}_i$, which is introduced in the next Section along the description of the data set. The joint effect of high inflation and uncertainty about future monetary policy is then modeled in terms of the interaction $d_{i,t-1}^{tolerant,\pi>\mathcal{T}_i} = d_{i,t-1}^{tolerant} \times d_{i,t-1}^{\pi>\mathcal{T}_i}$. Moreover, since the sample period in our empirical analysis covers several decades, it is likely that the threshold varies over time. Since the estimation of time- and economy-specific thresholds is likely inefficient, we employ a measure which quantifies temporary deviations of inflation from its long-term trajectory. This metric is given by $\tilde{\pi}_{i,t-1}^{GAP} = \pi_{i,t-1} - \bar{\pi}_{i,t-2}^{t-6}$, where $\bar{\pi}_{i,t-2}^{t-6} = (1/5)\sum_{j=1}^{5} \pi_{i,t-j-1}$ and is referred to as the inflation gap in the following. Measures which are similar to $\tilde{\pi}_{i,t-1}^{GAP}$ are employed by Cogley (2005) or Stock and Watson (2010) for deviations of inflation and unemployment, respectively, from their long-term trajectories. The corresponding interaction term with the effect of monetary policy uncertainty is given by $d_{i,t-1}^{tolerant} \times \tilde{\pi}_{i,t-1}^{GAP}$.

[5] In the following, a monetary authority's chairperson is referred to as "central bank governor", irrespective of whether the actual title is "governor", "chairman" or "president", etc.

[6] The sensitivity of the empirical analysis regarding this classification of $d_{i,t-1}^{tolerant}$ for Eurozone economies is examined in Section 5.5.

Finally, the error process of the SUR model is given by (9) and (10). The covariance matrix Σ is specified such that it allows for heteroscedasticity and nonzero correlations among the disturbances $(\epsilon_{1,t}, ..., \epsilon_{N,t})'$. This structure of the error term is taken into account by means of SUR estimation of the model described in (8) to (10). In the representation (9), unobservable influences on $IU_{i,t}$ are decomposed into a global time-fixed effect as denoted by λ_t on the one hand and country specific characteristics on the other hand. We think of the time-fixed effect λ_t mainly as representing the *Great moderation*. As in Engle and Rangel (2008), we separate cross section-specific characteristics into time-invariant country fixed effects η_i and short- to medium term dynamics. The latter are modeled via the AR specification of $v_{i,t}$ in (10). Time-invariant country specific effects may arise from distinct historical experiences such as periods of excess inflation, e.g. the German hyperinflation period during the years 1920-1923 (Alesina and Summers, 1993). A source of idiosyncratic dynamics in $IU_{i,t}$ might be (unexpected) FX rate adjustments or incidences of fiscal dominance (Davig et al., 2011). In the latter case, monetary policy decisions might be restricted during times of increasing government deficits.

4 Data

Our data set comprises annualized monthly CPI inflation, given by $\pi_{i,t,m} = 1200 \times \ln(CPI_{i,t,m}/CPI_{i,t,m-1})$, and the growth rate of the industrial production (IP) index denoted as $y_{i,t,m} = 1200 \times \ln(IP_{i,t,m}/IP_{i,t,m-1})$. All series are obtained from *Datastream* and seasonally adjusted by means of the X12 method. The cross section covers 13 advanced economies: Canada, Denmark, Finland, France, Germany, Italy, Norway, Portugal, Spain, Sweden, Switzerland, the UK and the U.S.[7] While the *annualized* monthly rates $\pi_{i,t,m}$ and $y_{i,t,m}$ are used to estimate the parameters of the Spline-GARCH model described in (1) to (2), the *annual* rates $\pi_{i,t} = 1/12 \sum_{m=1}^{12} \pi_{i,t,m}$ and $y_{i,t} = 1/12 \sum_{m=1}^{12} y_{i,t,m}$ are employed as explanatory variables in the subsequent analysis of low-frequency $IU_{i,t}$. The sample covers the period between 1975:1 and 2010:12. With $m = 1, ..., 12$ and $T = 36$, the dataset consists of 432 monthly observations. The output gap is given by $\tilde{y}_{i,t,m} = ip_{i,t,m} - ip^{HP}_{i,t,m}$, where $ip^{HP}_{i,t,m}$ is the long-term trend of $ip_{i,t,m} = \ln(IP_{i,t,m})$ as estimated by the Hodrick-Prescott filter with the smoothing parameter set to 129600, as it is suggested by Ravn and Uhlig (2002) for monthly data. Annual series $\tilde{y}_{i,t}$ obtain as $\tilde{y}_{i,t} = (1/12) \sum_{m \in t} \tilde{y}_{i,t,m}$. Table 1 reports summary statistics for the average yearly inflation rates, $\bar{\pi}_i = (1/T) \sum_t \pi_{i,t}$, their standard deviations, $\overline{SD}_i(\pi) = (1/T) \sum_t SD_{i,t}(\pi)$ and the corresponding statistics $\bar{y}_i$ and $\overline{SD}_i(y)$ which summarize the dynamics in $y_{i,t}$.

As can be seen from Table 1, $\bar{\pi}_i$ and also $\overline{SD}_i(\pi)$ vary considerably across countries. To take this heterogeneity into account, we define the country-specific threshold indicator variable $d^{\pi > \mathcal{T}_i}_{i,t} = \mathbb{1}\{\pi_{i,t} > \bar{\pi}_i + \overline{SD}_i(\pi)\}$, which equals unity in case of unusually high inflation rates. The indicator $d^{tolerant}_{i,t}$ is based on information regarding the timing of central bank governors' turnovers. The concept to connect the convictions of central bank governors to the political conditions at the time of their appointment is discussed, e.g., in Chapell et al. (1993) and is employed in empirical studies of distinct governments' influence on monetary policy by Grier (1991) or Belke and Potrafke (2012). The classification of governments is taken from data constructed in Beck et al. (2001). Appointment dates, in turn, are provided by Sturm and De Haan

[7] For the UK, we determine $\pi_{i,t,m}$ by employing the so-called "retail price index" which is the most widely used price index in this country.

Table 1: Country-specific summary statistics

	$\bar{\pi}_i$	$\overline{SD}_i(\pi)$	$\bar{y}_i$	$\overline{SD}_i(y)$
Canada	4.04	0.84	2.34	2.62
Denmark	4.22	0.99	1.98	19.82
Finland	4.41	0.82	3.00	12.10
France	4.19	0.57	0.93	9.46
Germany	2.39	0.67	1.58	5.41
Italy	6.46	0.56	1.12	9.83
Norway	4.56	0.95	2.65	14.27
Portugal	9.36	1.72	2.25	9.46
Spain	6.91	1.03	1.11	7.10
Sweden	4.61	1.15	1.86	7.47
Switzerland	2.10	0.75	0.79	0.68
UK	5.66	0.94	0.57	10.24
U.S.	4.02	0.68	2.19	2.15

Cell entries report averages $\bar{\pi}_i = (1/T)\sum_t \pi_{i,t}$ and $\overline{SD}_i(\pi) = (1/T)\sum_t SD_{i,t}(\pi)$ in columns 2 and 3, respectively. The statistics $\bar{y}_i$ and $\overline{SD}_i(y)$ are computed analogously.

(2001). This data set is also discussed in Dreher et al. (2008, 2010). The index $indep_i$ combines information regarding the appointment procedure and composition of the monetary policy committee, the influence of governments on monetary policy decision taking, the objectives and instruments of monetary policy and the regulations regarding financial support of national budgets through monetary policy. The values $indep_i$ can assume ranges from zero in case of no independence to a value of eight for complete independence (Sousa, 2002). Data on central banks' target rates $R_{i,t}$ are provided by the *International Monetary Fund*. Correlation statistics between $IU_{i,t}$ and its potential determinants which enter (8) as parts of $x_{i,t-1}$ and $D_{i,t-1}$ are summarized in Table 2.

The correlations in Table 2 show a strong relation between $IU_{i,t}$ and $\pi_{i,t}$. There is also a sizeable correlation between $IU_{i,t}$ and $SD_{i,t}(\pi)$, whereas $y_{i,t}$ and the corresponding volatility are less strongly correlated with $IU_{i,t}$. In contrast, measures of monetary

Table 2: Correlations between $IU_{i,t}$ and its potential determinants

	$IU_{i,t}$	$\pi_{i,t}$	$\tilde{\pi}_{i,t}^{GAP}$	$y_{i,t}$	$SD_{i,t}(\pi)$	$SD_{i,t}(y)$	$\lvert\tilde{R}_{i,t}\rvert$	$\tilde{R}_{i,t}^{+}$	$\tilde{R}_{i,t}^{-}$	$Vr_{i,t}(R)$
$\pi_{i,t}$	0.66	·								
$\tilde{\pi}_{i,t}^{GAP}$	0.34	0.29	·							
$y_{i,t}$	0.07	0.05	-0.20	·						
$SD_{i,t}(\pi)$	0.86	0.60	0.29	0.03	·					
$SD_{i,t}(y)$	0.11	0.14	0.03	-0.03	0.10	·				
$\lvert\tilde{R}_{i,t}\rvert$	0.65	0.78	0.32	0.03	0.59	0.11	·			
$\tilde{R}_{i,t}^{+}$	-0.12	-0.23	-0.21	-0.05	-0.12	0.06	0.07	·		
$\tilde{R}_{i,t}^{-}$	-0.67	-0.83	-0.31	-0.05	-0.61	-0.08	-0.91	0.33	·	
$Vr_{i,t}(R)$	0.18	0.23	0.01	-0.04	0.25	-0.01	0.20	0.21	-0.10	·
$d_{i,t}^{tolerant}$	0.04	-0.04	-0.02	-0.04	0.03	0.06	-0.05	-0.01	0.04	-0.01

policy like $|\tilde{R}_{i,t}|$, $\tilde{R}_{i,t}^{-}$ or $Vr_{i,t}(R)$ are more strongly related to $IU_{i,t}$. Moreover, we find only small correlations among most of the quantities in the last five rows of Table 2. This suggests that the distinct ways we employ to evaluate monetary policy might deliver independent information on the emergence of $IU_{i,t}$.

5 Empirical results

In this Section, we first summarize economy-specific diagnostics for the Spline-GARCH model outlined in equations (1) to (5). Second, we examine graphically the country-specific trajectories of the $IU_{i,t}$ series as implied by the estimates of the Spline-GARCH model. Third, the estimation results from the model in (8) and (9) are reported and discussed. Finally, we assess the robustness of the empirical findings with respect to model specification, alternative choices of the dependent variable and the sample period. We also compare the employed measure of $IU_{i,t}$ to a survey-based proxy of inflation uncertainty.

5.1 IU at the yearly frequency

Table 3 summarizes the estimation results from the Spline-GARCH model. In the second and third column, the lag orders $P_i^{(\pi)}$ and $P_i^{(y)}$ for the Phillips curve in (1) are reported. In all countries, the selected lag order as recommended by the BIC is higher for inflation than the one for output growth. The parameter estimates from the unity GARCH specification in (5) are given in columns 4 and 5. Moreover, the parameter K_i in the spline function in (4), which determines the long-term behavior of the $IU_{i,t}$-trajectories is shown in the rightmost column of Table 3. For all economies except Norway, the BIC suggests $K_i = 1$. This means that the estimation of $\tau_{i,t}$ in (4) is confined to the most slowly evolving fluctuations.

Table 3: Model specification diagnostics for (4) and (8)

	$P_i^{(\pi)}$	$P_i^{(y)}$	α_i	β_i	K_i
Canada	5	2	0.19	0.56	1
Denmark	9	0	0.11	0.09	1
Finland	7	0	0.11	0.89	1
France	8	0	0.14	0.62	1
Germany	10	4	0.26	0.51	1
Italy	6	0	0.28	0.23	1
Norway	8	0	0.18	0.28	2
Portugal	8	0	0.29	0.11	1
Spain	10	0	0.21	0.45	1
Sweden	8	0	0.38	0.12	1
Switzerland	6	1	0.06	0.84	1
UK	6	3	0.32	0.48	1
US	4	0	0.20	0.65	1

Note: The reported numbers are rounded to two decimals. Estimates for Finland satisfy $\alpha_i + \beta_i < 1$.

The graphs in Figure 1 display the country-specific evolution of $IU_{i,t}$. Although the plots show substantial differences in the evolution of $IU_{i,t}$ across the 13 economies, the countries can be broadly separated into two categories. France, Germany and

Switzerland are characterized by comparably low and stable levels of $IU_{i,t}$ during the whole sample period. For the remaining countries, the trajectories of $IU_{i,t}$ show a marked decline during the first half of the sample period. This remarkable similarity is usually referred to as the *Great moderation* (Blanchard and Simon, 2001; Benati, 2008) and is potentially the result of a rule-based and predictable monetary policy (Taylor, 2012). The dynamics of $IU_{i,t}$ varies across these economies mainly in terms of the magnitude of the reduction. However, $IU_{i,t}$ shows a tendency to increase again from the year 2000 onwards until the end of the sample in several economies. This increase of $IU_{i,t}$ is clearly visible for Canada, Norway the UK and the U.S. Less pronounced increases can be observed for the EMU member economies Portugal and Spain. The contrast between the more tranquil period during the *Great moderation* and the subsequent uprise of uncertainty is described in Taylor (2012) for the case of the U.S. Taylor (2012) associates the increase in uncertainty since the 2000's with failures of monetary policy to adhere to transparent and predictable rules. Though it refers to a more general macroeconomic context, the emphasis on the impact of policy uncertainty described by Taylor (2012) resembles the discussion of Ball (1992), where uncertainty about the conduct of future monetary policy is the main source of $IU_{i,t}$.

5.2 Macroeconomic influences on IU

In the following, the parameter estimates of the SUR model given by (8) and (9) are discussed. Specification I in Table 4 relates $IU_{i,t}$ solely to the macroeconomic quantities $\pi_{i,t-1}$, $y_{i,t-1}$ and their respective volatilities $SD_{i,t-1}(\pi)$ and $SD_{i,t-1}(y)$.[8] This first specification is intended to replicate the setup of previous studies which focus on the macroeconomic determinants of $IU_{i,t}$. In line with Grier and Perry (1998, 2000), Conrad and Karanasos (2005a,b) or Hartmann and Herwartz (2012), we find a positive and significant effect of $\pi_{i,t-1}$ on $IU_{i,t}$. The findings of Grier and Perry (1998, 2000) which refer to inflation uncertainty at high frequency are confirmed by our analysis of the low-frequency component of $IU_{i,t}$. Furthermore, we find that $IU_{i,t}$ is significantly and positively related to $SD_{i,t-1}(\pi)$. That is, uncertainty increases with the variability in inflation, which is, e.g., in line with empirical findings by Fischer (1981) for the U.S. The coefficient estimate reported in Table 4 shows that an increase in $SD_{i,t-1}(\pi)$ by one percentage point is accompanied by a higher $IU_{i,t}$ of about 0.26 percentage points. In stark contrast, neither $y_{i,t-1}$ nor its variability, $SD_{i,t-1}(y)$, appear to be significantly related to $IU_{i,t}$. This finding is line with Mankiw et al. (2003, p.229) who report that uncertainty (disagreement) shows "no clear relationship with measures of real activity". In addition, we obtain estimates of time-fixed effects λ_t from (9). The trajectory of $\hat{\lambda}_t$, which expresses the cross-sectional average trend in $IU_{i,t}$ is depicted in Figure 2. As for the case of the country-specific plots of $IU_{i,t}$, $\hat{\lambda}_t$ reflects the reduction of inflation uncertainty during the *Great moderation* period. Moreover, $\hat{\lambda}_t$ indicates that the cross-sectional average inflation uncertainty is increasing at the end of the sample period.

[8]Note that to increase legibility, the coefficient estimates in all Tables are multiplied by a factor of 100.

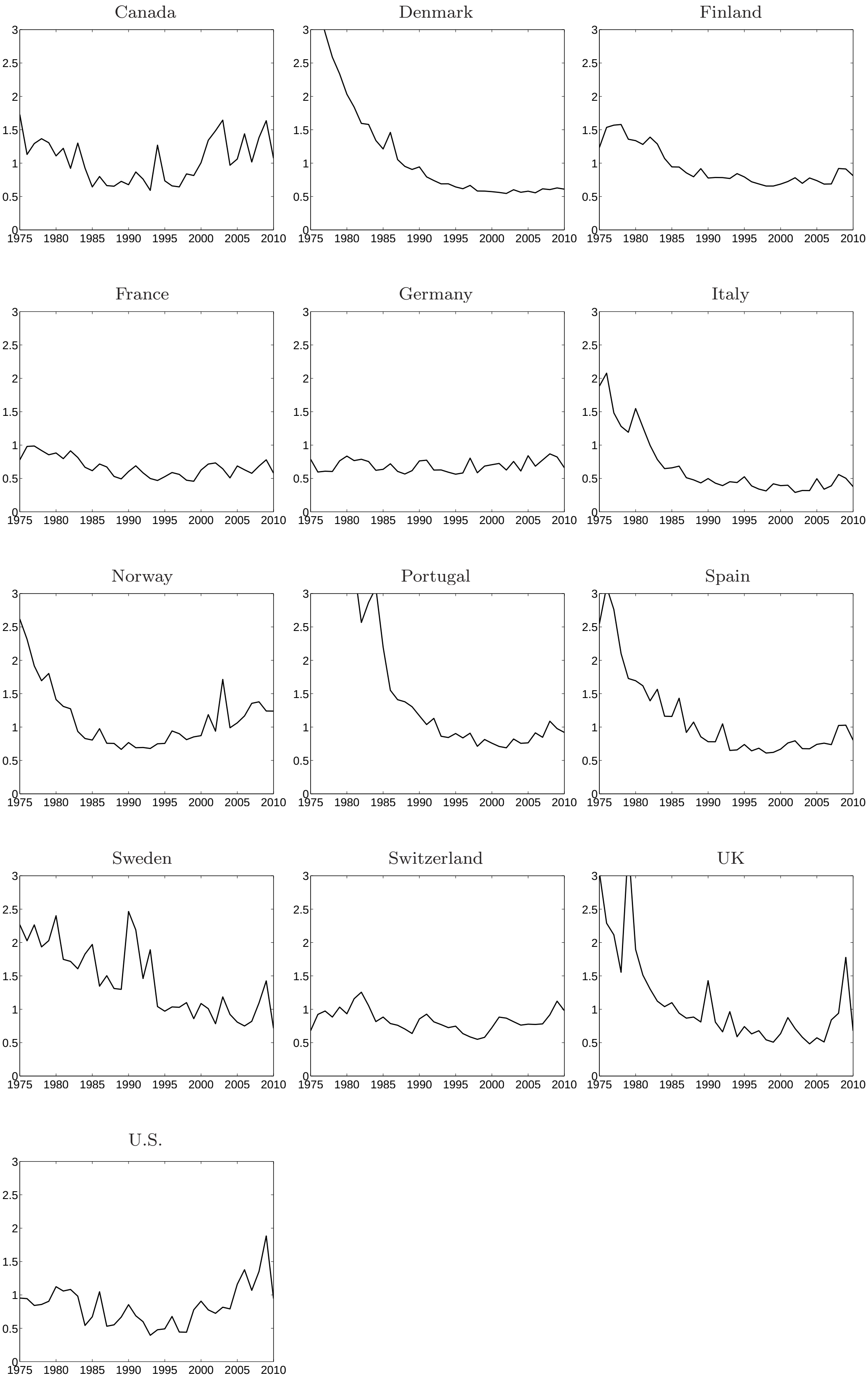

Figure 1: $IU_{i,t}$ from model (7) and (8) for 13 economies.

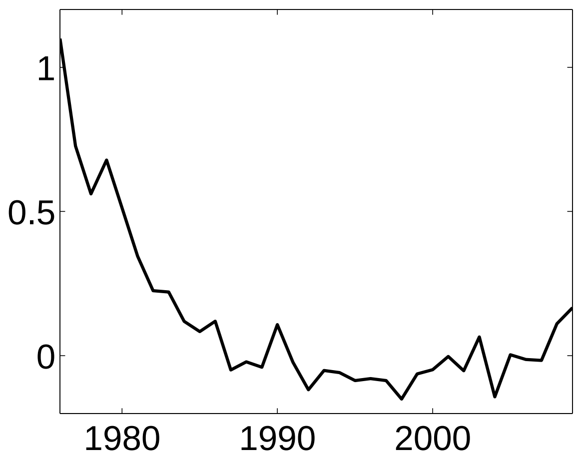

Figure 2: The long-term trend of $IU_{i,t}$ as denoted $\hat{\lambda}_t$ in (9).

5.3 The relation between monetary policy and IU

Next, specification I in Table 4 is extended by variables which summarize the characteristics of monetary policy. We particularly emphasize the relation between $IU_{i,t}$ and the characterization of monetary policy schemes as more or less inflation-averse on the one hand or ad-hoc versus rule-based on the other hand. The first set of estimates are reported in column II to V.

First, we examine how the degree of central bank independence, $indep_i$, relates to $IU_{i,t}$. The negative coefficient estimates show that $IU_{i,t}$ is lower for higher degrees of $indep_i$. Thus, economies where monetary policy is less affected by political influences are characterized by lower $IU_{i,t}$. Similarly, Alesina and Summers (1993) document for a cross section of 16 industrialized economies and the time period between 1973 and 1988 that countries with less independent central banks are characterized by less stable inflation rates. A relation between the disagreement of inflation expectations and central bank independence is also documented by Dovern et al. (2012).

Second, we evaluate monetary policy by means of $\tilde{R}_{i,t-1}$, the deviations of the realized policy rate from the value implied by the Taylor rule. As shown in column II to IV, the effect of the deviations' magnitude $|\tilde{R}_{i,t-1}|$ is positive and significant at the 5%-level, i.e. deviations from the Taylor rule path are associated with increasing $IU_{i,t}$. In specification V, $|\tilde{R}_{i,t-1}|$ is replaced by the factors $\tilde{R}^{+}_{i,t-1}$ and $\tilde{R}^{-}_{i,t-1}$ which quantify the effects of monetary expansions and contractions separately. The coefficient of $\tilde{R}^{-}_{i,t-1}$ is significant at the 5%-level, whereas $\tilde{R}^{+}_{i,t-1}$ is insignificant. The parameter estimate related to $\tilde{R}^{-}_{i,t-1}$ is negative, which means that $IU_{i,t}$ tends to be higher during periods when monetary policy is expansive beyond the degree which is recommended by the Taylor rule. Interestingly, the significant coefficient of $\pi_{i,t-1}$ in column II turns insignificant if measures based on $\tilde{R}_{i,t-1}$ are included (columns III to V)[9]. The significance of $\pi_{i,t-1}$ in specification II could be rationalized by regarding inflation as a crude proxy of an inadequately loose monetary policy. The relatively high correlation between $\pi_{i,t-1}$ and $|\tilde{R}_{i,t-1}|$ as shown in Table 2 underlines this suggestion.

Third, $IU_{i,t}$ might be influenced by the variation in short-term interest rates, $Vr_{i,t-1}(R)$. Dovern et al. (2012) examine the impact of a metric similar to $Vr_{i,t-1}(R)$ on the disagreement of inflation expectations and find that disagreement increases during periods of strong interest rate fluctuations. However, in our case, the linkage between $IU_{i,t}$ and $Vr_{i,t-1}(R)$ is insignificant across all model reformulations reported in Table 4. This finding does not change if, e.g., covariates based on $\tilde{R}_{i,t-1}$ are excluded. It is also evident from the correlation statistics in Table 2 that $Vr_{i,t-1}(R)$ is not strongly related to $|\tilde{R}_{i,t-1}|$, $\tilde{R}^{+}_{i,t-1}$ or $\tilde{R}^{-}_{i,t-1}$, thus the explanatory content of $Vr_{i,t-1}(R)$ seems to be low.

[9]The significance of $\pi_{i,t-1}$ seems to be unaffected if only $indep_i$ is considered in addition to the explanatory variables in column I.

Fourth, our model includes the dummy variable $d_{i,t-1}^{tolerant}$ which indicates whether in country i in year $t-1$ a less inflation averse governor is in power and the (country specific) inflation threshold dummy $d_{i,t-1}^{\pi>\mathcal{T}_i}$. Comparing columns II and III of Table 4 shows that the coefficient estimates on $\pi_{i,t-1}$, $y_{i,t-1}$, $SD_{i,t-1}(\pi)$, $SD_{i,t-1}(y)$, $indep_i$ and $|\tilde{R}_{i,t-1}|$ remain almost unchanged after the additional variables are included. Moreover, the estimated coefficients of both $d_{i,t-1}^{tolerant}$ and $d_{i,t-1}^{\pi>\mathcal{T}_i}$ are positive but only $d_{i,t-1}^{tolerant}$ is significant. That is, $IU_{i,t}$ appears to be higher during the time when less inflation-averse governors are in power. However, monetary policy schemes which put less emphasis on low inflation might increase $IU_{i,t}$ in particular during periods of high inflation. This mechanism is described in detail in Ball (1992). We examine this hypothesis in the specifications IV and V which include the interaction term $d_{i,t-1}^{tolerant,\mathcal{T}_i}$. We allow for an economy-specific threshold level since it is possible that the influence of monetary policy comes into effect at different levels of $\pi_{i,t-1}$ for distinct economies. This could, for example, be due to different historical experiences with periods of high and uncertain inflation. The coefficient estimate for $d_{i,t-1}^{tolerant,\mathcal{T}_i}$ shows that $IU_{i,t}$ is significantly higher if $\pi_{i,t-1} > \mathcal{T}_i$ and $d_{i,t-1}^{tolerant} = 1$. This clearly confirms that prediction of Ball's (1992) model. The size of the estimated effect of $d_{i,t-1}^{tolerant,\mathcal{T}_i}$ is also considerably larger than the one for $d_{i,t-1}^{tolerant}$ and $d_{i,t-1}^{\pi>\mathcal{T}_i}$ alone in specification III. Hence, this influence on $IU_{i,t}$ comes into effect mainly if higher inflation rates prevail during the mandate of less inflation-averse governors.

In these specifications, $IU_{i,t}$ may increase if governors who tolerate higher inflation are in office and inflation raises above a specific threshold value. As an alternative, we account for potential time-variation in this threshold by reestimating the specifications I to V after the interaction term $d_{i,t-1}^{tolerant,\mathcal{T}_i}$ is replaced by $d_{i,t-1}^{tolerant} \times \tilde{\pi}_{i,t-1}^{GAP}$. The corresponding

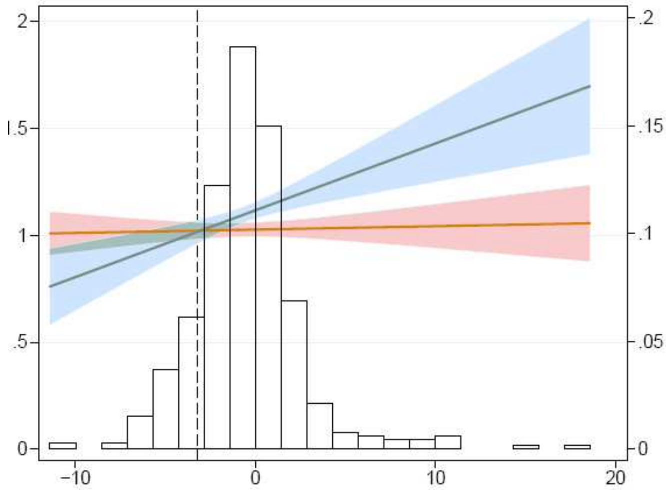

Figure 3: Predictions $\hat{IU}_{i,t}$ for $d^{tolerant}_{i,t-1} = 1$ (blue line) and $d^{tolerant}_{i,t-1} = 0$ (red line). Shaded areas lines depict the 95%-confidence intervals. The intersection of both predictions is indicated by the vertical line. Below, a histogram of $\tilde{\pi}^{GAP}_{i,t-1}$ is depicted.

estimates are summarized in columns VI to X of Table 4. Specification VI yields results which are relatively similar to the ones from model I. In contrast to column II and III, however, $\tilde{\pi}^{GAP}_{i,t-1}$ remains significant if $indep_i$ and $|\tilde{R}_{i,t-1}|$ are incorporated (columns VII and VIII). Apart from most other estimates which are basically similar in the specifications I to V on the one hand and VI to X on the other hand, the indicator $d^{tolerant}_{i,t-1}$ remains significant after $d^{tolerant}_{i,t-1} \times \tilde{\pi}^{GAP}_{i,t-1}$ is included, which is not the case in specifications IV and V. Our estimation results imply that – independently of the level of $\tilde{\pi}^{GAP}_{i,t-1}$ – a deviation of inflation from its trend leads to a stronger increase in $IU_{i,t}$ when the central bank governor is perceived as less inclined to adopt a disinflationary monetary policy ($d^{tolerant}_{i,t-1} = 1$). Figure 3 illustrates this result graphically. It shows a comparison of predicted values, $\hat{IU}_{i,t}$, given $d^{tolerant}_{i,t-1} = 0$ (red line) or $d^{tolerant}_{i,t-1} = 1$ (blue line) as a function of the level of the lagged inflation gap. The predictions of $\hat{IU}_{i,t}$ are obtained by setting all covariates in (8) except $\tilde{\pi}^{GAP}_{i,t-1}$ and the country- and time-fixed effects in (9) to their average values. The plots show that the slope of the relation between $IU_{i,t}$ and $\tilde{\pi}^{GAP}_{i,t-1}$ is steeper for the case when a central bank governor is appointed under a liberal government. The vertical line indicates the level of $\tilde{\pi}^{GAP}_{i,t-1}$ where the linear predictions $\hat{IU}_{i,t}$ intersect. If $\tilde{\pi}^{GAP}_{i,t-1}$ increases beyond a level of about -3%, the initially lower value of $\hat{IU}_{i,t}$ for $d^{tolerant}_{i,t-1} = 1$ exceeds the prediction given $d^{tolerant}_{i,t-1} = 0$. Put differently, the separation of inflation-tolerant from inflation-averse monetary policy schemes is increasingly important for the explanation of $IU_{i,t}$ if inflation rises above its long-term path. Moreover, shaded lines represent 95%-confidence intervals. The histogram in Figure 3 shows that the the intersection point is located below the majority of observations regarding $\tilde{\pi}^{GAP}_{i,t-1}$. Furthermore, the share of cases where $d^{tolerant}_{i,t-1} = 0$ to the total number of cases is almost equal in both parts of the sample.

5.4 Alternative volatility measures

In this Section, we consider two alternative proxies for the unobservable inflation uncertainty as dependent variables in the SUR estimation. A first natural candidate is the *ex-post* inflation variability measure $SD_{i,t}(\pi)$. As Engle et al. (2012) and Conrad and Loch (2013) point out, the *ex-post* measure $SD_{i,t}(\pi)$ can be considered a noisy proxy of the *ex-ante* measure $IU_{i,t}$. This argument is graphically illustrated by Figure 4 which depicts the trajectories of both $IU_{i,t}$ and $SD_{i,t}(\pi)$ over the sample period for the case of the U.S. The column labeled $SD_{i,t}(\pi)$ in Table 5 presents parameter estimates when $IU_{i,t}$ is replaced by $SD_{i,t}(\pi)$ as dependent variable. To facilitate comparability the column labeled $IU_{i,t}$ reproduces the parameter estimates of column III of Table 4, where $IU_{i,t}$ is the dependent variable. Similar results obtain regarding the influence of $indep_i$ and $\tilde{R}^{-}_{i,t-1}$, where, in particular, the impact of $indep_i$ reinstates the ones from Alesina and Summers (1993). The effect of lagged inflation variability is also significant and larger than the one for $IU_{i,t}$ in the leftmost column, which is not surprising given the high degree of persistence in inflation variability. In contrast to the case for $IU_{i,t}$, however, there is a significant influence of $SD_{i,t}(y)$ on $SD_{i,t}(\pi)$. In addition, neither lagged inflation nor $d_{i,t-1}^{tolerant}$ and $d_{i,t-1}^{tolerant,\pi>\mathcal{T}_i}$ have a significant effect on $SD_{i,t}(\pi)$. These findings highlight the importance of distinguishing between *ex-ante* and *ex-post* measures when analyzing the determinants of long-term inflation uncertainty. The findings for the case where $\bullet = \tilde{\pi}_{i,t-1}^{GAP}$ are rather similar to those for $\bullet = \pi_{i,t-1}$. However, the effect of $d_{i,t-1}^{tolerant,\pi>\mathcal{T}_i}$ is significant in this specification, possibly as a result of a stronger relation of the more transitory fluctuations which characterize $\tilde{\pi}_{i,t-1}^{GAP}$ to $SD_{i,t}(\pi)$.

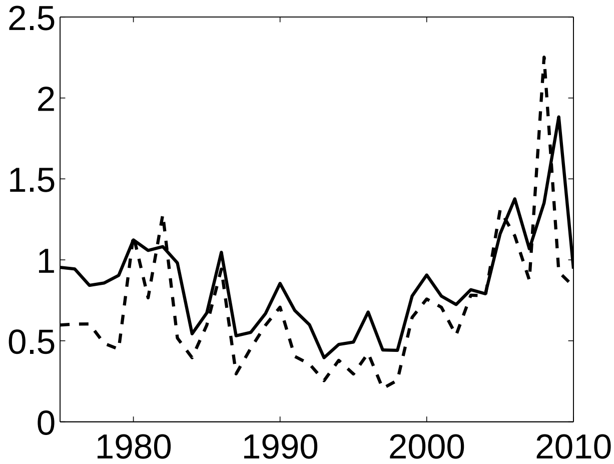

Figure 4: U.S. inflation uncertainty as measured by $IU_{i,t}$ (solid line) and $SD_{i,t}(\pi)$ (dashed line).

The second measure we consider comes from a standard GARCH model which assumes that the unconditional (country specific) variance of inflation is constant over time. Assuming that $\tau_{i,t} = \tau_i$, equation (6) reduces to

$$\widetilde{IU}_{i,t} = \tau_i^{1/2} \left(\sum_{m \in t} g_{i,t,m} \right)^{1/2} . \tag{13}$$

Although the unconditional variance is constant, $\widetilde{IU}_{i,t}$ varies from year-to-year since $\sum_{m \in t} g_{i,t,m}$ can be low in certain years but high in others depending on the size of the inflation forecast errors. Figure 5 shows the dynamics of $\widetilde{IU}_{i,t}$ in the U.S. at the annual frequency. As Table 5 reports, using $\widetilde{IU}_{i,t}$ as dependent variable, we still find that central bank independence, periods of overly expansive monetary policy and inflation-tolerant governors in times of high inflation lead to more inflation uncertainty. Although the changes in $\widetilde{IU}_{i,t}$ are now entirely driven by variation in the short-term component, this result might be explained by the fact that $\widetilde{IU}_{i,t}$ still extracts some long-term information by aggregating the $g_{i,t,m}$ over the year. In contrast, approaches that would directly rely on the monthly $g_{i,t,m}$ series are obviously much less suited to measure the long-term effects of changes in the conduct of monetary policy.

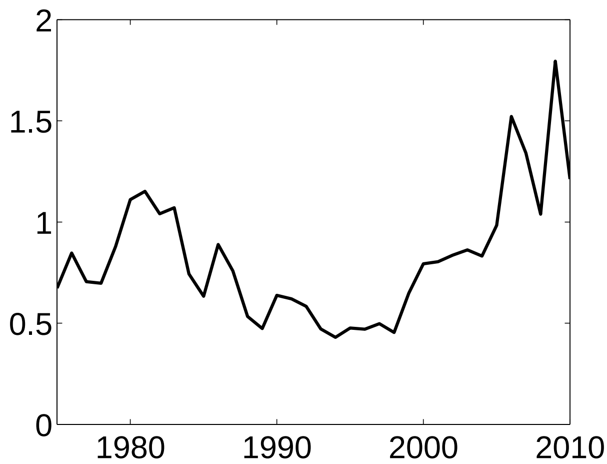

Figure 5: $\widetilde{IU}_{i,t}$ as measured by a standard GARCH model for the U.S.

5.5 Robustness analysis

A first objection against our findings might be that they are driven by the years after the unfolding of the financial- and sovereign debt crisis in 2008. It is possible that during this period, $IU_{i,t}$ has been higher due to increased uncertainty about the economic outlook in general. In order to show that this does not affect our findings regarding the relation between monetary policy and $IU_{i,t}$, we reestimate our model for a sample which does not include the years after 2008. In Table 6, the corresponding parameters estimates can be found in the column labeled "before 2008". Clearly, our findings are robust to excluding the most recent observations, the coefficient estimates hardly differ from those in column V of Table 4.

A second way in which the empirical findings might be distorted is through potential mis-classification of the EMU monetary policy regime, since 6 out of 13 economies in our cross section have delegated their responsibility for monetary policy to the ECB after the formation of the EMU. Thus, we report estimates for a reduced cross-section which excludes all EMU member states. Moreover, as described in Section 4, for the EMU economies we have coded $d_t^{tolerant} = 0$ after the inception of the Euro. We choose this specification of $d_t^{tolerant}$ since it has been argued that the way how the ECB conducts monetary policy can be related to the policy decisions of the (presumably rather inflation averse) German Bundesbank prior to the EMU period (Hayo and Hofmann, 2006). However, it is also possible that the monetary policy of the ECB has become less affected by the Bundesbank but more influenced by EMU economies with a higher preference for discretionary monetary policy. For example, Sturm and Wollmershäuser (2008) find that small member countries have relatively high voting power in monetary policy decisions. Moreover, Faust et al. (2001) find that an estimated reaction function of the ECB reveals a higher emphasis on output stabilization than the corresponding Bundesbank estimate. Thus, in an alternative specification we set $d_{i,t-1}^{tolerant} = 1$ instead of $d_{i,t-1}^{tolerant} = 0$ for the Euro zone economies since the advent of the EMU. These estimates are reported in the second column of Table 6. As the results in Table 6 show, our main findings hold in both specifications.

A further robustness check could be to analyze $IU_{i,t}$ and $\pi_{i,t}$ jointly in a bivariate specification. Whereas Friedman (1977) or Ball (1992) describe the influence of $\pi_{i,t}$ on $IU_{i,t}$, Cukierman and Meltzer (1986) argue that $IU_{i,t}$ might, in turn, affect $\pi_{i,t}$ if certain monetary policy strategies are adopted. However, empirical examinations of the causal impact of $IU_{i,t}$ on $\pi_{i,t}$ document limited evidence in favor of the latter hypothesis. In particular, multi-economy studies find that the sign and significance of the impact of $IU_{i,t}$ on $\pi_{i,t}$ differs across industrialized economies (Daal et al., 2005; Hartmann and Herwartz, 2012). Thus, we abstain from specifying the relation between $IU_{i,t}$ and $\pi_{i,t}$ by means of a bivariate system.

5.6 Comparison of $IU_{i,t}$ with survey based inflation uncertainty

As discussed in Section 1, ideally we would measure inflation uncertainty directly using survey data. However, survey expectations on inflation are only available for selected countries and restricted time periods. Nevertheless, in this Section we compare our model based uncertainty measure for the U.S., $IU_{US,t}$, with data from the FED's Survey of Professional Forecasters. Following Capistrán and Ramos-Francia (2010) or Dovern et al. (2012) we use the disagreement among forecasters as a measure of inflation uncertainty. Specifically, we measure disagreement, $Dis_{US,t}$, by the interquartile range of the individual forecasters one-year-ahead point predictions of the annualized quarterly growth rate of the CPI. We choose this forecast horizon since the suitability of disagreement as a proxy of uncertainty deteriorates especially at longer anticipation periods (Lahiri and Sheng, 2010). Figure 6 depicts the evolution of $Dis_{US,t}$ and $IU_{i,t}$. The evolution of the two series is apparently similar. In both cases, we observe the downward trend associated with the *Great moderation* since the early 1980s and raising levels of uncertainty towards to end of the sample period. Nevertheless, the graph makes also clear that $IU_{i,t}$ is leading with respect to $Dis_{US,t}$ which is confirmed by a simple cross-correlation analysis. This leading property may be due to the fact that the spline function in equation (4) essentially defines a two-sided filter which employs forward-looking information to determine current long-term uncertainty.

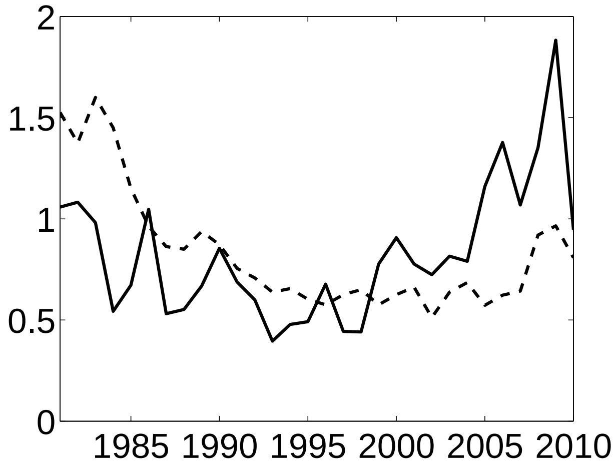

Figure 6: U.S. inflation uncertainty as measured by $IU_{US,t}$ (Spline-GARCH, solid line) and $Dis_{US,t}$ (interquartile range of inflation forecasts from the Survey of Professional Forecasters, dashed line).

6 Conclusions

We analyze the determinants of long-term inflation uncertainty for 13 industrialized economies. Long-term inflation uncertainty is measured as a time-varying unconditional variance in the framework of the Spline-GARCH model as introduced by Engle and Rangel (2008).

We find that monetary policy and macroeconomic conditions do not act as independent sources of inflation uncertainty but that the most sizeable increases in inflation uncertainty occur if less inflation-averse central bank governors are in charge of monetary policy during periods when inflation is high. Following Ball (1992), our interpretation of this finding is that governors' attitudes can increase uncertainty when the question if inflation will finally be reduced becomes particularly relevant during high-inflation periods.

A further significant effect stems from deviations of target interest rates from the path prescribed by the Taylor rule (Taylor, 1993). This underlines the role of ad-hoc monetary policy schemes as a source of inflation uncertainty. In particular, inflation uncertainty increases after actual interest rates have been lower than the level as suggested by the Taylor rule, i.e. after periods of unduely expansive monetary policy.

Moreover, central bank independence as a further aspect of the monetary policy framework is significantly related to inflation uncertainty. In line with previous empirical evidence on the relation between inflation and monetary policy independence (Alesina and Summers, 1993), find that economies with more independent central bank are characterized by lower inflation uncertainty. In contrast, we do not find that the variability of the target interest rate matters, i.e. frequent adjustments of interest rates are not associated with increasing inflation uncertainty.

An examination of alternative approximations of inflation uncertainty shows that the Spline-GARCH-implied metric is more suitable than other approaches such as the intra-annual standard deviation of inflation or the annualized conditional variance from a conventional GARCH(1,1) model. The former are ex-post metrics and fail to detect the theoretically asserted joint influence of monetary policy and inflation dynamics on inflation uncertainty. The latter disregard time-variation of the low-frequency component of inflation uncertainty and are therefore incomplete measures for the analysis of the implications of monetary policy and macroeconomic conditions which likely unfold their impact over longer time periods.

The documented effects are robust with respect to subsample choices and a variety of model reformulations. Excluding the sample observations from the year 2008 onwards shows that the high aggregate uncertainty does not qualitative affect our conclusions. Similarly, the classification of monetary policy schemes as ad-hoc or rule-based, which is more difficult for the Eurozone member states than for single economies, is not a crucial driver of the reported outcomes. Excluding all members of the European Monetary Union which are part of the cross section (6 out of 13) leads to the same findings as considering the entire sample.

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Table 4: Coefficient estimates ×100 from (8)

Model number:	I	II	III	IV	V	VI	VII	VIII	IX	X
	$\bullet = \pi_{i,t-1}$					$\bullet = \tilde{\pi}^{GAP}_{i,t-1}$				
$\bullet$	**1.35**	0.55	0.36	0.24	−1.19	**1.34**	**1.80**	**1.72**	0.08	0.16
	(4.37)	(1.36)	(0.74)	(0.48)	(−1.70)	(4.07)	(5.00)	(4.57)	(0.17)	(0.34)
$y_{i,t-1}$	0.09	0.10	0.15	0.08	−0.03	0.11	0.14	0.15	0.15	0.10
	(0.60)	(0.68)	(0.99)	(0.50)	(−0.21)	(0.73)	(0.95)	(1.02)	(0.87)	(0.59)
$SD_{i,t-1}(\pi)$	**25.78**	**24.62**	**23.15**	**22.31**	**21.37**	**23.61**	**21.38**	**20.61**	**23.89**	**22.80**
	(11.45)	(10.93)	(9.89)	(9.72)	(9.27)	(10.41)	(9.34)	(8.83)	(9.41)	(8.80)
$SD_{i,t-1}(y)$	0.02	−0.13	−0.13	0.08	0.19	0.08	−0.04	−0.08	−0.02	−0.03
	(0.14)	(−0.89)	(−0.83)	(0.52)	(1.12)	(0.53)	(−0.23)	(−0.52)	(−0.09)	(−0.14)
$indep_i$		**−7.52**	**−6.57**	**−7.98**	**−10.24**		**−8.20**	**−7.14**	−2.33	−2.72
		(−3.62)	(−3.08)	(−3.32)	(−4.04)		(−3.42)	(−2.92)	(−0.86)	(−1.04)
$\lvert\tilde{R}_{i,t-1}\rvert$		**1.25**	**1.20**	**1.26**			**1.28**	**1.26**	**1.04**	
		(3.94)	(3.63)	(3.67)			(4.56)	(4.33)	(3.12)	
$\tilde{R}^{+}_{i,t-1}$					−0.15					0.48
					(−0.24)					(0.73)
$\tilde{R}^{-}_{i,t-1}$					**−2.45**					**−1.45**
					(−5.04)					(−3.89)
$SD_{i,t-1}(R)$		−0.30	−0.48	−0.51	−0.12		−0.11	−0.18	−0.15	−0.06
		(−0.59)	(−0.89)	(−0.95)	(−0.21)		(−0.21)	(−0.32)	(−0.24)	(−0.09)
$d^{tolerant}_{i,t-1}$			**3.25**	0.45	0.16			3.18	**8.74**	**9.00**
			(2.00)	(0.25)	(0.09)			(1.80)	(4.23)	(4.28)
$d^{\pi_i > \mathcal{T}_i}_{i,t-1}$			3.64	−1.32	−1.23					
			(1.28)	(−0.38)	(−0.34)					
$d^{tolerant,\mathcal{T}_i}_{i,t-1}$				**18.47**	**19.69**					
				(4.89)	(5.03)					
$d^{tolerant}_{i,t-1} \times \tilde{\pi}^{GAP}_{i,t-1}$									**2.99**	**2.97**
									(3.79)	(3.72)

Note: Roman numbers represent alternative specifications of (8). Cell entries in boldface indicate coefficient estimates which are significant at the 5%-level. Below the coefficients, t-statistics are reported in parentheses. Deviations of inflation from its long-run trend are computed as $\tilde{\pi}^{GAP}_{i,t-1} = \pi_{i,t} - \bar{\pi}^{t-5}_{i,t-1}$, with $\bar{\pi}^{t-5}_{i,t} = (1/5)\sum_{j=1}^{5} \pi_{i,t-j}$.

Table 5: Coefficient estimates ×100 from (8), alternative dependent variables

Dependent variable:	$SD_{i,t}(\pi)$	$\widetilde{IU}_{i,t}$	$SD_{i,t}(\pi)$	$\widetilde{IU}_{i,t}$
	$\bullet = \pi_{i,t-1}$		$\bullet = \tilde{\pi}^{GAP}_{i,t-1}$	
$\bullet$	−1.01	−0.03	0.39	0.54
	(−1.02)	(−0.04)	(0.62)	(1.15)
$y_{i,t-1}$	0.03	0.01	0.36	0.04
	(0.09)	(0.06)	(1.26)	(0.24)
$SD_{i,t-1}(\pi)$	**31.62**	**34.70**	**29.80**	**32.30**
	(7.51)	(16.24)	(7.01)	(12.84)
$SD_{i,t-1}(y)$	**0.78**	0.19	**0.79**	0.08
	(2.84)	(1.02)	(2.74)	(0.42)
$indep_i$	**−11.05**	**−8.05**	**−8.04**	−5.87
	(−3.17)	(−3.10)	(−2.61)	(−1.92)
$\tilde{R}^{+}_{i,t-1}$	0.30	0.62	0.86	0.53
	(0.30)	(1.08)	(0.90)	(0.80)
$\tilde{R}^{-}_{i,t-1}$	**−3.79**	**−3.21**	**−3.31**	**−2.79**
	(−4.84)	(−6.69)	(−6.48)	(−6.82)
$Vr_{i,t-1}(R)$	0.16	0.69	−0.16	0.70
	(0.15)	(1.14)	(−0.14)	(0.94)
$d^{tolerant}_{i,t-1}$	−0.32	0.48	4.34	**6.14**
	(−0.11)	(0.28)	(1.67)	(3.11)
$d^{\pi>\mathcal{T}_i}_{i,t-1}$	3.05	−1.57		
	(0.53)	(−0.46)		
$d^{tolerant,\mathcal{T}_i}_{i,t-1}$	6.13	**16.28**		
	(1.04)	(4.47)		
$d^{tolerant}_{i,t-1} \times \tilde{\pi}^{GAP}_{i,t-1}$			**3.66**	**1.61**
			(3.31)	(1.97)

For a description, see Table 4.

Table 6: Coefficient estimates ×100 from (8): Robustness analysis

Sample:	Before 2008	Non-Euro	$d_{i,t-1}^{tolerant} = 1$ f. Eurozone
$\pi_{i,t-1}$	-0.98	0.42	-0.89
	(-1.48)	(0.28)	(-1.27)
$y_{i,t-1}$	-0.09	-0.19	-0.04
	(-0.53)	(-0.54)	(-0.27)
$SD_{i,t-1}(\pi)$	**19.42**	**22.83**	**21.83**
	(8.47)	(5.97)	(9.50)
$SD_{i,t-1}(y)$	0.24	0.42	0.19
	(1.37)	(1.19)	(1.12)
$indep_i$	**$-$9.81**	-66.73	**$-$10.01**
	(-3.80)	(-1.41)	(-3.98)
$\tilde{R}_{i,t-1}^{+}$	0.36	-1.74	-0.03
	(0.62)	(-1.25)	(-0.05)
$\tilde{R}_{i,t-1}^{-}$	**$-$2.61**	-1.09	**$-$2.33**
	(-5.82)	(-0.99)	(-4.82)
$Vr_{i,t-1}(R)$	-0.33	**4.73**	-0.14
	(-0.64)	(3.63)	(-0.24)
$d_{i,t-1}^{tolerant}$	1.14	-6.31	-2.18
	(0.64)	(-1.58)	(-1.20)
$d_{i,t-1}^{\pi > \mathcal{T}_i}$	-1.73	-6.18	-3.54
	(-0.52)	(-0.93)	(-0.98)
$d_{i,t-1}^{tolerant,\mathcal{T}_i}$	**22.30**	**24.16**	**22.09**
	(6.32)	(3.21)	(5.75)

Notes: Results obtained by setting $d_{i,t-1}^{tolerant} = 1$ instead of $d_{i,t-1}^{tolerant} = 0$ for the Eurozone economies after 1999 are stated in the rightmost column. For further descriptions, see Table 4.