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

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Abstract

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 long-term variance component of inflation in the framework of the 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 and if the conduct of monetary policy is ad-hoc rather than rule-based.

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1 Introduction

The impact of uncertainty shocks on macroeconomic performance has been intensively debated in the recent literature (see, e.g., Bloom, 2009; Fernández-Villaverde et al., 2011; and Bloom et al., 2014). Specifically, the interaction between monetary policy and uncertainty has received considerable attention (Taylor, 2012; Bekaert et al., 2013). One type of uncertainty that has been long recognized to have negative welfare effects is inflation uncertainty (IU) (see, e.g., Fischer and Modigliani, 1978; Barnea et al., 1979). In particular, Friedman (1977) and, more recently, Taylor (2012) discuss how certain types of monetary policy can give rise to IU or macroeconomic uncertainty in general. In a theoretical model, Ball (1992) formalizes the idea of Friedman (1977) and emphasizes that it is particularly the interaction of monetary policy and inflation which generates IU. Moreover, Ball and Cecchetti (1990) stress that the relation between changes in monetary policy and IU should be most clearly detectable when considering long-term rather than short-term IU.

In this study, we consider a range of policy measures that assess the stance of monetary policy. The relation of these metrics to IU is evaluated in terms of a general multi-country model for 13 developed economies which allows for the interaction 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 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. Specifying IU in terms of the conditional variance of the unpredictable component of inflation is in the spirit of the measure of time-varying macroeconomic uncertainty recently suggested in Jurado et al. (2014).[1]

We do not consider survey-based measures of IU, which are an alternative approach that is commonly adopted in the related literature (Zarnowitz and Lambros, 1987; Giordani and Söderlind, 2003). This is because such data is available only for short time periods and for a limited set of economies such as the Euro area or the US 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 different monetary policy schemes are based on the quantification of deviations from the Taylor rule (Taylor 1993) and the appointment dates of central bank governors.[2] 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, we distinguish central bank governors by their degree of inflation-aversion. According to Nordhaus (1975), Samuelson (1977), Alesina and Sachs (1988) or Berger and Woitek (2005), liberal governors are typically perceived as more inflation-tolerant. Hence, we follow Sturm and De Haan (2001) and Dreher et al. (2008, 2010) and relate changes in a monetary authority's degree of inflation-tolerance or -aversion to the appointment dates of central bank governors. 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.

We analyze the relation between IU and its potential determinants in an empirical model that allows for cross-country dependencies and unobserved characteristics of the IU process which vary across economies and time periods. The estimation of a cross-

[1]In contrast to the IU measure we consider, the index proposed in Jurado et al. (2014) refers to overall economic uncertainty in the US and is based on 132 macroeconomic indicators.

[2]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.

sectional average trend in IU shows that long-term IU has been decreasing during the *Great Moderation* period until the year 2003, but subsequently rising since the unfolding of the recent financial- and sovereign debt crisis until the end of the sample period in the year 2010.

Our main findings can be summarized as follows. First, we show 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. In particular, IU increases if inflation-tolerant governors are in power during high-inflation periods. This confirms the theoretical argument of Ball (1992) that high inflation leads to high IU if there is uncertainty about the central bank's willingness to disinflate. This joint effect of inflation and the preferences of monetary policy is markedly stronger than the influence of inflation in isolation, which is a commonly adopted means of explaining IU in the empirical literature.

Second, 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. This finding is in line with the argument put forth in Taylor (2012) that unpredictable, i.e. ad-hoc rather than rule-based, monetary policy creates uncertainty and, thereby, leads to poor economic performance. Interestingly, the Bank of International Settlements (BIS) reports that for several countries – which are included in our study – market expectations of future policy rates are currently below the trajectory prescribed by the Taylor rule and argues that the "the risk of normalizing too late and too gradually should not be underestimated" (BIS, 2014, p.101). Our findings suggest increasing IU as one of the channels through which such risks could materialize.

By considering alternative 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 IU. As expected, pure measures of ex-post inflation variability, which are often associated with IU, are less suitable since such metrics can only be regarded as noisy approximations of ex-ante uncertainty. Importantly, by quantifying IU with an ex-post measure such as the intra-yearly variability of inflation, we obtain results which indicate a weak or seemingly missing relation between IU, changes in the conduct of monetary policy and the interaction of monetary policy with the level of inflation. We also document that our measure of long-term IU and the interquartile range of inflation expectations from the Survey of Professional Forecasters of the US FED are strongly related.

Additionally, we provide several robustness checks and show that our main results remain unaffected. Among other things, we document that IU increases with the volatility in global equity markets which compliments the findings in Engle and Rangel (2008) and Conrad and Loch (2014). In addition, our results show that IU is lower in countries where the central bank is legally declared as being independent.

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. 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 of IU is the level of inflation. Widely cited discussions of the relationship between inflation and IU include Okun (1971), Friedman (1977), Fischer and Modigliani (1978) or Cukierman and Meltzer (1986). Ball (1992) formalizes the hypothesis of a causal impact of inflation on IU. 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 the level of 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 the possibility that distinct monetary policy schemes have different effects. Moreover, the interaction of macroeconomic conditions and the monetary policy framework is typically disregarded.

Several studies examine the influence of different monetary policy schemes on IU. In a study on the relation between inflation and IU in the US, Evans and Wachtel (1993) document that changes in the monetary policy regime are an important determinant of IU and argue that such regime changes occur only infrequently. They argue that changes in 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. 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 proxy IU by the root mean squared error (RMSE) computed from a cross section of survey-based inflation expectations and find that IU tends to be higher under more inflation-tolerant regimes.

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 Bank of England adopts a formal inflation target. Caporale and Kontonikas (2009) show that the relation betwen IU and inflation in European economies is affected by the formation of the European Monetary Union (EMU) in the year 1999. Similarly, Hartmann and Herwartz (2013) document that IU is significantly smaller in EMU economies after the introduction of the Euro 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) report 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.

Though the influence of macroeconomic conditions and the characteristics of monetary policy on IU has been documented in several studies, these determinants are typically considered one at a time, thereby effectively disregarding potentially important interaction effects. Such an interaction effect is described by the theoretical model of Ball (1992), where the combination of higher levels of inflation and uncertainty about the degree of inflation tolerance of monetary policy drives IU.

3 Measuring and analyzing long-term IU

IU is an unobservable quantity. The choice as to which of the alternative proxies that have been proposed to measure it is most suitable depends on the question under consideration. 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 Phillips 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).[3] 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 for 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

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

which allows for secular variations in IU. Specifically, Ball and Cecchetti (1990) model inflation as a random walk plus noise and show that the effect of higher levels of inflation on IU becomes more pronounced at longer horizons.[4] Similarly, Fischer (1981) argues that the way in which monetary policy is implemented can be expected primarily to affect the low-frequency movements in IU. Following these arguments, we employ a measure of long-term or low-frequency IU in this study. The measure will be based on the Spline-GARCH model of Engle and Rangel (2008) which allows us to separate the conditional variance of inflation into a short-term and a long-term component, with the latter changing smoothly over time.

We adopt a two-stage procedure similar to the one employed by Engle and Rangel (2008). First, we estimate the coefficients of the Spline-GARCH model based on monthly observations and then aggregate the conditional variances to a yearly frequency. Second, the implied low-frequency (yearly) IU measure is related to indicators of institutional conditions and economic quantities for which only annual observations are available.

3.1 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 production is denoted as $y_{i,t,m}$. 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 $y_{i,t,m}$ are denoted $P_i^{(\pi)}$ and $P_i^{(y)}$,

[4]The model used in Ball and Cecchetti (1990) is observationally equivalent to an IMA(1,1). Stock and Watson (2007) show that this specification adequately describes the US inflation process.

respectively, and are selected by the BIC. The maximum lag order is set to twelve.[5] 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, for instance.[6]

Modeling the Phillips curve relation based on industrial production 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).[7]

We assume that the innovations to inflation are given by

$$\begin{aligned} u_{i,t,m} &= \sqrt{h_{i,t,m}} Z_{i,t,m}, \quad Z_{i,t,m} \overset{iid}{\sim} (0,1), && (2) \\ \text{with } \; h_{i,t,m} &= \tau_{i,t} g_{i,t,m}, && (3) \end{aligned}$$

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 a monthly frequency and is intended to capture the transitory component of inflation volatility, $\tau_{i,t}$ changes at a yearly frequency only and reflects long-term influences such as changes in the institutional conditions of monetary policy. The long-term trend in IU is modeled as an exponential spline function given by

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

In (4), the flexibility of the trend function increases with the order K_i. Short-term IU is

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

[6] For example, our notation should be understood as $\pi_{i,t,0} = \pi_{i,t-1,M}$.

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

expressed in terms of a unit variance GARCH process, which reads as

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

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 $\mathbf{E}[u^2_{i,t,m}] = \mathbf{E}[g_{i,t,m}\tau_{i,t}Z^2_{i,t,m}] = \tau_{i,t}$, 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, whereby the BIC guides the selection of K_i.

We base our measure of IU on the monthly series of conditional variances $\mathbf{E}_{i,t,m-1}[u^2_{i,t,m}] = h_{i,t,m}$, 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 an annual frequency. For this, we define our measure of annual IU in country i and year t as the square root of the aggregated monthly conditional variances:

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

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, 2014). In analogy to (7), we calculate the annual variability of output, denoted by $SD_{i,t}(y)$.

3.2 Determinants of IU

Apart from uncertainty regarding future monetary policy and macroeconomic factors, $IU_{i,t}$ can be driven by various other factors that are mostly outside the range of decision making at the national level. The importance of such factors is reflected in the debate over 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:

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

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

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

and $(\epsilon_{1,t}, ..., \epsilon_{N,t})' \overset{iid}{\sim} (0, \mathbf{\Sigma})$. In (8), the predetermined macroeconomic quantities are summarized in $\mathbf{x}_{i,t-1} = (\pi_{i,t-1}, y_{i,t-1}, SD_{i,t-1}(\pi), SD_{i,t-1}(y))'$. 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 drivers 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 $SD_{i,t-1}(y)$.

Next, we introduce several metrics which quantify the influence of monetary policy conditions on $IU_{i,t}$. These determinants are summarized in the vector $\mathbf{D}_{i,t-1}$. Two measures in $\mathbf{D}_{i,t-1}$ are based on the interest rate, $R_{i,t-1}$, which is set by a country's central bank. The first measure is derived from the *Taylor rule*, a widely used means to quantify the predictability of monetary policy. Following (Taylor, 1993), 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- and long-run value, respectively. Though not all economies in the cross section have explicitly announced inflation targets, during recent decades inflation targeting has become a widespread monetary policy rule among industrialized economies. Leaving aside a particular recommendation about the most suitable values for γ_π and γ_y, we measure the extent to which monetary policy corresponds with the original specification of Taylor (1993), in which $\gamma_\pi = 1.5$, $\gamma_y = 0.5$ and a level of 2% for both r_i and $\pi_i^\star$ is assumed.[8] The same specification is also employed, for example, by Nikolsko-Rzhevskyy et al. (2014) or the BIS to compare market-implied interest rates to the ones prescribed by the Taylor rule (BIS, 2014). Similarly, we use the absolute value of past deviations of the actual interest rate from the target rate suggested by the Taylor rule, denoted as $|\tilde{R}_{i,t-1}|$, where $\tilde{R}_{i,t-1} = R_{i,t-1} - R^\star_{i,t}$, as a potential determinant of IU. However, the response of $IU_{i,t}$ to contractionary (positive) and expansionary (negative) deviations is not necessarily symmetric. Thus, in an alternative specification, $\mathbf{D}_{i,t-1}$ 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.[9] Moreover, the variability of short-term interest rates might reflect a lack of smoothness in the way how monetary policy is conducted. Based on quarterly interest rates $R_{i,t-1,q}$, we compute a measure for the steadiness of the interest rates set by a central bank 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 variability of interest rates and distinct forms of inertia in central banks' behavior. We consider metrics

[8] We also considered forward-looking Taylor rule specifications as proposed by Clarida et al. (1998), where the parameters have been estimated by means of the generalized method of moments. Respective results are qualitatively similar to those based on (11) and will be provided by the authors upon request.

[9] Alternatively, a binary distinction between rule-based and ad-hoc ways to conduct monetary policy is obtained by means of identifying periods where the most pronounced deviations from a particular prescription for monetary policy occur. For example, Nikolsko-Rzhevskyy et al. (2013) adopt such a strategy by detecting structural breaks in Taylor rule deviations.

such as $\tilde{R}_{i,t-1}$ and $Vr_{i,t-1}(R)$ as quantitative measures of the monetary policy stance.[10]

Moreover, monetary policy regimes may be classified as rather inflation-tolerant or inflation-averse. We separate these two monetary policy schemes by means of the dummy variable $d_{i,t-1}^{tolerant}$, which is one if a central bank governor was appointed under a left-wing government and zero in all other cases.[11] 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). Our specification can be thought of as an implementation of the theoretical model in Ball (1992), where liberal central bank governors give rise to $IU_{i,t}$ because, in contrast to other governors, they cannot be expected to disinflate during high-inflation periods. In this model, $IU_{i,t}$ is triggered by the appointment of inflation-tolerant central bank governors only if inflation exceeds a certain level. Thus, in addition to $d_{i,t-1}^{tolerant}$, $\mathbf{D}_{i,t-1}$ 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 $\mathcal{T}_i$. The selection procedure for the threshold $\mathcal{T}_i$ is data-driven and will be 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 via the interaction term $d_{i,t-1}^{tolerant,\pi>\mathcal{T}_i} = d_{i,t-1}^{tolerant} \times d_{i,t-1}^{\pi>\mathcal{T}_i}$. Moreover, since our sample period covers several decades, the country-specific threshold may also vary over time. Since the estimation of time- and economy-specific thresholds is likely to be 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 (2002) or Stock and Watson (2010) for deviations of inflation

[10] Note that a high value of $Vr_{i,t-1}(R)$ does not necessarily imply that monetary policy is not predictable. High values of $Vr_{i,t-1}(R)$ can also occur if monetary policy is rule-based, but adjusts to rapidly changing macroeconomic conditions.

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

and unemployment, respectively, from their long-term trajectories. The corresponding interaction term with the type of central bank governor is given by $d^{tolerant}_{i,t-1} \times \tilde{\pi}^{GAP}_{i,t-1}$.

Finally, the error process of the SUR model is given by (9) and (10). The covariance matrix $\mathbf{\Sigma}$ 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 denoted by λ_t on the one hand and country specific characteristics on the other hand. We mainly think of the time-fixed effect λ_t 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 episodes 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) exchange 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 covers a cross section of $N = 13$ advanced economies: Canada, Denmark, Finland, France, Germany, Italy, Norway, Portugal, Spain, Sweden, Switzerland, the UK and the US. All series (except interest rates) are obtained from *Datastream* and seasonally adjusted by means of the X12 method. The data on interest rates set by central banks, $R_{i,t}$, are provided by the *International Monetary Fund*.[12] Annualized monthly CPI inflation is calculated as $\pi_{i,t,m} = 1200 \times \ln(CPI_{i,t,m}/CPI_{i,t,m-1})$ and the growth rate of the industrial

[12]We thank Matthias Neuenkirch for sharing the dataset in a readily useable format with us.

production (IP) index as $y_{i,t,m} = 1200 \times \ln(IP_{i,t,m}/IP_{i,t,m-1})$.[13] The sample covers the period between 1975:1 and 2010:12. With $m = 1, ..., 12$ and $T = 36$, the dataset consists of 13×432 monthly observations. 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 analysis of low-frequency $IU_{i,t}$.

For each country, Table 1 reports the average yearly inflation rate, $\bar{\pi}_i = (1/T)\sum_t \pi_{i,t}$, the average of the intra-yearly standard deviations, $\overline{SD}_i(\pi) = (1/T)\sum_t SD_{i,t}(\pi)$, and the corresponding statistics $\bar{y}_i$ and $\overline{SD}_i(y)$ for industrial production. As can be seen from Table 1, both $\bar{\pi}_i$ and $\overline{SD}_i(\pi)$ vary considerably across countries. To take this heterogeneity into account, we define the country-specific threshold indicator variable $d_{i,t}^{\pi>\mathcal{T}_i} = \mathbb{1}\{\pi_{i,t} > \bar{\pi}_i + \overline{SD}_i(\pi)\}$, which equals unity in case of "high" inflation rates.

[Place Table 1 here]

The indicator $d_{i,t}^{tolerant}$ distinguishes between inflation-averse and -tolerant central bank governors. If a governor is appointed during the term of a liberal government, we code $d_{i,t}^{tolerant} = 1$, whereas $d_{i,t}^{tolerant} = 0$ in all other cases. The classification of governments is taken from data constructed in Beck et al. (2001), who distinguish between right-wing, left-wing and centrist governments by assessing the respective governing party's stance towards economic policy. Appointment dates, in turn, are provided by Sturm and De Haan (2001). This data set is also discussed in Dreher et al. (2008, 2010).

Finally, the output gap is given by $\tilde{y}_{i,t,m} = ip_{i,t,m} - ip_{i,t,m}^{HP}$, where $ip_{i,t,m}^{HP}$ 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 suggested by Ravn and Uhlig (2002) for monthly data. Annual series $\tilde{y}_{i,t}$ are obtained as $\tilde{y}_{i,t} = (1/12)\sum_{m\in t} \tilde{y}_{i,t,m}$.

[13] 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.

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 graphically examine the country-specific trajectories of the $IU_{i,t}$-series as implied by the estimates of the Spline-GARCH model and provide correlation statistics between our measure of $IU_{i,t}$ and its potential determinants. Third, the estimation results for 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 2 summarizes the estimation results for 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. The lag orders selected for inflation are between 4 and 10, whereas according to the BIC industrial production is only relevant in four (with lag orders between 1 and 4) out of the thirteen economies. The parameter estimates for the unit variance GARCH specification in (5) are given in columns 4 and 5 and imply that the short-term component is covariance stationary for all countries. Moreover, the parameter K_i in the spline function in (4) is shown in the rightmost column of Table 2. 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.

[Place Table 2 here]

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, for several economies $IU_{i,t}$ shows a tendency to increase from the year 2000 onwards. This rise of $IU_{i,t}$ is clearly visible for Canada, Norway, the UK and the US. 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 discussed in Taylor (2012) for the case of the US. Taylor (2012) associates the increase in uncertainty beginning in 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 $IU_{i,t}$.

[Place Figure 1 here]

Table 3 presents the averages of the correlations between $IU_{i,t}$ and its potential determinants within the 13 economies. As expected, the correlation statistics show a strong relation between $IU_{i,t}$ and $\pi_{i,t}$. Unsurprisingly, there is an even stronger correlation between $IU_{i,t}$ and $SD_{i,t}(\pi)$, whereas $y_{i,t}$ and the corresponding volatility are only weakly correlated with $IU_{i,t}$. In contrast, the deviations from the Taylor rule, in particular $\tilde{R}^{-}_{i,t}$, are strongly related to $IU_{i,t}$. The correlation between both $Vr_{i,t}(R)$ and $d^{tolerant}_{i,t}$ with $IU_{i,t}$ is rather low. Moreover, the mutual correlations among the measures that are based on the deviations from the Taylor rule, $Vr_{i,t}(R)$ and $d^{tolerant}_{i,t}$ are comparably low. This suggests that the distinct ways we employ to evaluate monetary policy might deliver independent information on the emergence of $IU_{i,t}$.

[Place Table 3 here]

5.2 IU and the macroeconomy

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 $SD_{i,t-1}(y)$.[14] This first specification is intended to replicate the findings of previous studies that 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 $IU_{i,t}$.[15] 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. 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 inflation uncertainty (disagreement) shows "no clear relationship with measures of real activity".

Figure 2 shows the trajectory of the estimated time-fixed effect $\hat{\lambda}_t$ in (9), i.e. the cross-sectional time trend in $IU_{i,t}$. 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 towards the end of the sample period.

[Place Figure 2 here]

[14] To increase readability, the coefficient estimates in all Tables are multiplied by a factor of 100.

[15] In contrast to our approach, previous studies have mainly focused on individual economies and monthly measures of IU.

[Place Table 4 here]

5.3 The relation between monetary policy and IU

Next, specification I in Table 4 is extended by including variables that reflect the stance 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 and ad-hoc versus rule-based on the other hand. The first set of estimates are reported in columns II to V.

First, 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 columns 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 are associated with increasing $IU_{i,t}$. In specification V, $|\tilde{R}_{i,t-1}|$ is split into $\tilde{R}^{+}_{i,t-1}$ and $\tilde{R}^{-}_{i,t-1}$, which quantify the effects of overly expansionary and contractionary monetary policy schemes separately. Importantly, the coefficient on $\tilde{R}^{-}_{i,t-1}$ is negative and significant at the 5%-level which means that $IU_{i,t}$ tends to be higher when monetary policy is expansive beyond the degree which is recommended by the Taylor rule. In sharp contrast, the parameter estimate related to $\tilde{R}^{+}_{i,t-1}$ is insignificant, i.e. contractionary policies do not affect IU. Interestingly, the coefficient on $\pi_{i,t-1}$ turns insignificant if measures based on $\tilde{R}_{i,t-1}$ are included in columns II to V. Thus, the significance of $\pi_{i,t-1}$ in specification I 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 3 underlines this suggestion.

Second, $IU_{i,t}$ 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 highly volatile interest rates. 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 covariates based on $\tilde{R}_{i,t-1}$ are excluded. It is also evident from the correlation statistics in Table 3 that $Vr_{i,t-1}(R)$ is only weakly 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.

Third, column III of Table 4 presents parameter estimates for the case that the dummy variables $d^{tolerant}_{i,t-1}$ and $d^{\pi>\mathcal{T}_i}_{i,t-1}$ are included. The estimated coefficients of both $d^{tolerant}_{i,t-1}$ and $d^{\pi>\mathcal{T}_i}_{i,t-1}$ are positive but only $d^{tolerant}_{i,t-1}$ is significant. That is, $IU_{i,t}$ appears to be higher during episodes when less inflation-averse governors are in power.[16] A comparison of columns II and III shows that the coefficient estimates on $\pi_{i,t-1}$, $y_{i,t-1}$, $SD_{i,t-1}(\pi)$, $SD_{i,t-1}(y)$ and $|\tilde{R}_{i,t-1}|$ remain almost unchanged.

However, as discussed in Ball (1992), monetary policy schemes which put less emphasis on low inflation might increase $IU_{i,t}$ in particular during periods of high inflation. We examine this hypothesis in specifications IV and V which include the interaction term $d^{tolerant,\mathcal{T}_i}_{i,t-1}$. 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. The coefficient estimate for $d^{tolerant,\mathcal{T}_i}_{i,t-1}$ shows that $IU_{i,t}$ is significantly higher if $\pi_{i,t-1} > \mathcal{T}_i$ and $d^{tolerant}_{i,t-1} = 1$. This clearly confirms that prediction of Ball's (1992) model, i.e. this influence on $IU_{i,t}$ comes into effect mainly if higher inflation rates prevail during the mandate of more inflation-tolerant governors.

Next, we reestimate all models by replacing inflation as well as the country-specific threshold with the inflation gap variable. The corresponding estimates are summarized in columns VI to X of Table 4. Specification VI yields results which are similar to the ones from model I. In contrast to column II and III, however, $\tilde{\pi}^{GAP}_{i,t-1}$ remains significant if $|\tilde{R}_{i,t-1}|$ and $d^{tolerant}_{i,t-1}$ are incorporated (columns VII and VIII). Most importantly, in columns IX

[16] This finding is also in line with the theoretical results derived in Conrad and Eife (2012). Using a simple New Keynesian model, they show that inflation persistence as well as the variability of inflation increase (decrease) if a central bank places less (more) weight on inflation relative to output growth.

and X, the indicator $d_{i,t-1}^{tolerant}$ remains significant even after including $d_{i,t-1}^{tolerant} \times \tilde{\pi}_{i,t-1}^{GAP}$, which is not the case in specifications IV and V. Our estimation results imply that 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 ($d_{i,t-1}^{tolerant} = 1$). Figure 3 illustrates this result graphically. It shows a comparison of the predicted inflation uncertainties, $\hat{IU}_{i,t}$, given $d_{i,t-1}^{tolerant} = 0$ (solid red line) or $d_{i,t-1}^{tolerant} = 1$ (dashed blue line) as a function of the level of the lagged inflation gap. The predictions $\hat{IU}_{i,t}$ are obtained by setting all covariates in (8) except $\tilde{\pi}_{i,t-1}^{GAP}$ and the country- and time-fixed effects in (9) to their average values. The figure reveals that the predictions $\hat{IU}_{i,t}$ are flat in the case of conservative governors, i.e. under inflation-averse governors IU does not respond to changes in the inflation gap. In sharp contrast, for inflation-tolerant governors there is a positive relation between the inflation gap and IU. Figure 3 also shows the histogram of the inflation gap. Obviously, the effect on IU of being a more or less inflation-averse governor becomes more pronounced with larger inflation gaps.

[Place Figure 3 here]

5.4 Alternative volatility measures

In this Section, we consider two alternative measures for the unobservable inflation uncertainty as dependent variables in the SUR estimation. A first natural candidate is the *ex-post* inflation variability $SD_{i,t}(\pi)$. As Engle et al. (2013) and Conrad and Loch (2014) point out, the *ex-post* measure $SD_{i,t}(\pi)$ can be considered a noisy proxy for the *ex-ante* measure $IU_{i,t}$. The columns labeled $SD_{i,t}(\pi)$ in Table 5 present parameter estimates when $IU_{i,t}$ is replaced by $SD_{i,t}(\pi)$ as the dependent variable. Clearly, $SD_{i,t}(\pi)$ is strongly related to its own past lag and also to the variability in industrial production. While $\tilde{R}_{i,t-1}^{-}$ does have a strong impact on $SD_{i,t}(\pi)$, the link between $SD_{i,t}(\pi)$ and $d_{i,t-1}^{tolerant}$ appears to exist only when the regression is based on the inflation gap. These findings suggest that it can be important to distinguish between *ex-ante* and *ex-post* measures when analyzing the determinants of long-term inflation uncertainty.

[Place Table 5 here]

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, $\widetilde{IU}_{i,t}$ still 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. As Table 5 shows, using $\widetilde{IU}_{i,t}$ as the dependent variable, we again find that IU is driven by periods of overly expansive monetary policy and inflation-tolerant governors that are in power in times of high inflation. Although the changes in $\widetilde{IU}_{i,t}$ are now entirely driven by variation in the short-term component, our result can 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.

5.5 Robustness analysis

In this Section, we demonstrate the robustness of our empirical findings with respect to reformulations of the model described in (8) and (9).

First, we consider two additional explanatory variables for $IU_{i,t}$. As argued by, e.g., Conrad and Loch (2014), $IU_{i,t}$ may be related to fluctuations in financial markets. We proxy global stock market volatility by computing the (yearly) realized volatility measure

$$RV_{t-1} = \sqrt{\sum_{d \in t-1} r_{d,t-1}^2},$$

where $r_{d,t-1}$ denotes the daily return on the MSCI World Equity Index. The results in the second column of Table 6 show that the influence of RV_{t-1} on $IU_{i,t}$ is significantly positive and also comparably large. Further, economic policymakers outside a central bank might affect $IU_{i,t}$, especially if the mandate of monetary policy is not legally guaranteed to be independent of the government. Hence, we relate $IU_{i,t}$ to an indicator of central bank independence, denoted as $indep_{i,t-1}$, which equals unity during the years when the central bank of economy i is officially independent (see Dovern et al., 2012). The negative coefficient estimate in the third column of Table 6 shows that $IU_{i,t}$ is lower for higher degrees of $indep_{i,t-1}$. Thus, economies where monetary policy is less affected by political influences are characterized by lower $IU_{i,t}$. Our finding is in line with with Alesina and Summers (1993) who first established that countries with less independent central banks are characterized as having more volatile inflation rates. A relation between central bank independence and $IU_{i,t}$ or the disagreement of inflation expectations is also documented by Grier and Perry (1998) and Dovern et al. (2012), respectively.

Second, our findings might be distorted by observations from the years after the unfolding of the financial- and sovereign debt crisis in 2008. It is possible that during this period, $IU_{i,t}$ is higher due to increased uncertainty about the economic outlook in general. In order to show that this does not affect our findings, we reestimate our model for a sample that does not include the years after 2008. In Table 6, the corresponding parameter estimates can be found in the column labeled "before 2008". Clearly, our findings are robust to excluding the most recent observations.

The third robustness check uses an alternative method to specify common effects in the specification of the error term. Instead of estimating time-fixed effects as in (9), one may include cross-section averages $\bar{\mathbf{x}}_{t-1} = (1/N) \sum_{i=1}^{N} \mathbf{x}_{i,t-1}$ to proxy for unobserved effects common to all economies (Pesaran, 2006). The corresponding results are reported in the column labeled "with $\bar{\mathbf{x}}_{t-1},\ \lambda_t = 0$" and show that the conclusions drawn from column V of Table 4 remain valid if common disturbances are modeled in terms of $\bar{\mathbf{x}}_{t-1}$.

Fourth, the empirical findings might be affected through potential mis-classification of the EMU monetary policy regime, because 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. So far, we have coded $d_t^{tolerant} = 0$ for the EMU economies after the inception of the Euro. We choose this specification of $d_t^{tolerant}$ since it can be argued that the way the ECB has been set up is in the spirit of the (presumably rather inflation-averse) German Bundesbank (Hayo and Hofmann, 2006). However, it is also possible that the (recent) monetary policy of the ECB is 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 strong 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 Eurozone economies. The corresponding estimates are reported in the rightmost column of Table 6, which shows that our main findings are not affected. The interaction term now even takes a slightly higher value than before which might suggest that the ECB's policy is indeed best characterized as less inflation-averse.

[Place Table 6 here]

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

As discussed in Section 1, ideally we would like to 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 US, $IU_{US,t}$, 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 to measure IU. 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 it matches our focus on yearly IU and, in addition, it is well known that the suitability of disagreement as a proxy of uncertainty deteriorates with the forecast horizon (Lahiri and Sheng, 2010). Figure 4 depicts the evolution of $Dis_{US,t}$ and $IU_{i,t}$ which appear to be quite similar. In both cases, we observe the downward trend associated with the *Great Moderation* since the early 1980s and raising levels of uncertainty towards the end of the sample period. Nevertheless, the graph also makes it clear that $IU_{i,t}$ is leading with respect to $Dis_{US,t}$ which is further confirmed by a simple cross-correlation analysis. This leading property may be due to the fact that the Spline-GARCH model is not estimated in real-time and, therefore, essentially employs forward-looking information to determine current long-term uncertainty.

[Place Figure 4 here]

6 Conclusions

We analyze the determinants of long-term IU for 13 industrialized economies. Long-term IU is measured as the aggregated yearly conditional 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 IU 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 of high inflation. Following Ball (1992), our interpretation of this finding is that governors' attitudes towards inflation are an important driver of long-term IU.

A further significant effect materializes if interest rates set by central banks deviate from those prescribed by the Taylor rule (Taylor, 1993). This underlines the role of ad-hoc monetary policy schemes as a source of IU. In particular, IU increases after actual interest rates have been lower than the level suggested by the Taylor rule, i.e. after periods of unduely expansive monetary policy. Since global monetary policy can be currently described as very accommodative, our results suggest that one of the risks of “exiting too late or too slowly” could be increasing IU (BIS, 2014, p.99).

An examination of alternative approximations of IU 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 documented effects are robust with respect to restricting the sample period and a variety of model reformulations. Excluding the observations from the year 2008 onwards shows that our conclusions are not driven by the exceptionally high aggregate uncertainty during that period. Similarly, the classification of the monetary policy scheme of the ECB as more or less inflation-tolerant, which is more difficult than for single economies, is not a crucial driver of the reported outcomes. Moreover, we find that the volatility in international stock markets has a significant influence on IU. Furthermore, alternative ways of modeling dependencies in the disturbance process among the cross-section units leads to essentially identical findings.

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A Tables and Figures

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
US	4.02	0.68	2.19	2.15

Note: 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.

Table 2: Specification diagnostics for the Spline-GARCH model in (4)

	$P_i^{(\pi)}$	$P_i^{(y)}$	$\hat{\alpha}_i$	$\hat{\beta}_i$	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 $\hat{\alpha}_i + \hat{\beta}_i < 1$.

Table 3: Correlations between $IU_{i,t}$ and its potential determinants

	$IU_{i,t}$	$\pi_{i,t}$	$\tilde{\pi}_{i,t}^{GAP}$	$y_{i,t}$	$SD_{i,t}(\pi)$	$SD_{i,t}(y)$	$\lvert\tilde{R}_{i,t}\rvert$	$\tilde{R}_{i,t}^{+}$	$\tilde{R}_{i,t}^{-}$	$Vr_{i,t}(R)$
$\pi_{i,t}$	0.60	·								
$\tilde{\pi}_{i,t}^{GAP}$	0.33	0.34	·							
$y_{i,t}$	-0.01	0.05	-0.20	·						
$SD_{i,t}(\pi)$	0.77	0.48	0.32	-0.07	·					
$SD_{i,t}(y)$	0.13	0.15	0.08	-0.09	0.17	·				
$\lvert\tilde{R}_{i,t}\rvert$	0.56	0.69	0.26	0.02	0.44	0.12	·			
$\tilde{R}_{i,t}^{+}$	-0.19	-0.34	-0.21	-0.08	-0.15	0.02	0.11	·		
$\tilde{R}_{i,t}^{-}$	-0.61	-0.82	-0.34	-0.07	-0.47	-0.09	-0.85	0.39	·	
$Vr_{i,t}(R)$	0.18	0.13	0.05	-0.07	0.31	0.10	0.17	0.19	-0.07	·
$d_{i,t}^{tolerant}$	-0.08	-0.02	-0.05	-0.03	-0.07	-0.03	-0.04	0.04	0.03	-0.03

Note: For each country we first calculate the correlation between the different variables, e.g., $IU_{i,t}$ and $\pi_{i,t}$. The numbers reported are the averages of these correlations across the 13 countries.

Table 4: Macroeconomic and monetary policy-related determinants of $IU_{i,t}$ in the SUR model (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**	0.62	0.47	0.36	-0.81	**1.34**	**1.61**	**1.52**	-0.02	0.05
	(4.37)	(1.51)	(0.95)	(0.73)	(-1.19)	(4.07)	(4.38)	(3.92)	(-0.05)	(0.11)
$y_{i,t-1}$	0.09	0.09	0.15	0.08	-0.01	0.11	0.12	0.14	0.15	0.11
	(0.60)	(0.62)	(1.01)	(0.49)	(-0.05)	(0.73)	(0.81)	(0.92)	(0.85)	(0.61)
$SD_{i,t-1}(\pi)$	**25.78**	**24.69**	**22.98**	**22.48**	**21.51**	**23.61**	**21.42**	**20.50**	**23.95**	**22.77**
	(11.45)	(10.83)	(9.68)	(9.66)	(9.21)	(10.41)	(9.22)	(8.66)	(9.43)	(8.81)
$SD_{i,t-1}(y)$	0.02	-0.05	-0.06	0.14	0.24	0.08	0.03	-0.03	-0.00	-0.01
	(0.14)	(-0.38)	(-0.39)	(0.91)	(1.43)	(0.53)	(0.19)	(-0.16)	(-0.02)	(-0.06)
$\lvert\tilde{R}_{i,t-1}\rvert$		**1.20**	**1.12**	**1.18**			**1.18**	**1.15**	**1.03**	
		(3.68)	(3.31)	(3.46)			(4.09)	(3.82)	(3.09)	
$\tilde{R}^{+}_{i,t-1}$					-0.01					0.50
					(-0.02)					(0.76)
$\tilde{R}^{-}_{i,t-1}$					$\mathbf{-2.13}$					$\mathbf{-1.39}$
					(-4.44)					(-3.76)
$Vr_{i,t-1}(R)$		-0.16	-0.33	-0.46	-0.09		0.04	-0.06	-0.09	0.02
		(-0.30)	(-0.61)	(-0.87)	(-0.15)		(0.06)	(-0.11)	(0.65)	(0.03)
$d^{tolerant}_{i,t-1}$			**3.96**	1.32	1.38			**3.59**	**9.07**	**9.40**
			(2.40)	(0.73)	(0.74)			(1.99)	(2.05)	(4.51)
$d^{\pi_i > \mathcal{T}_i}_{i,t-1}$			2.94	-1.90	-1.80					
			(1.02)	(-0.56)	(-0.51)					
$d^{tolerant,\mathcal{T}_i}_{i,t-1}$				**17.76**	**18.16**					
				(4.80)	(4.76)					
$d^{tolerant}_{i,t-1} \times \tilde{\pi}^{GAP}_{i,t-1}$									**3.16**	**3.17**
									(4.01)	(3.97)

Note: Roman numbers represent alternative specifications of (8). Results are reported as coefficient estimates $\times 100$. 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: Results for alternative dependent variables

Dependent variable:	$SD_{i,t}(\pi)$		$\widetilde{IU}_{i,t}$	
	$\bullet = \pi_{i,t-1}$	$\bullet = \tilde{\pi}^{GAP}_{i,t-1}$	$\bullet = \pi_{i,t-1}$	$\bullet = \tilde{\pi}^{GAP}_{i,t-1}$
$\bullet$	-0.51	-0.00	0.75	0.22
	(-0.52)	(-0.00)	(1.24)	(0.46)
$y_{i,t-1}$	0.07	0.33	0.03	0.03
	(0.24)	(1.15)	(0.17)	(0.16)
$SD_{i,t-1}(\pi)$	**32.95**	**30.38**	**34.46**	**32.04**
	(7.74)	(7.09)	(16.01)	(12.60)
$SD_{i,t-1}(y)$	**0.78**	**0.77**	0.28	0.14
	(2.91)	(2.74)	(1.60)	(0.69)
$\tilde{R}^{+}_{i,t-1}$	-0.08	0.56	0.67	0.48
	(-0.08)	(0.58)	(1.19)	(0.72)
$\tilde{R}^{-}_{i,t-1}$	$\mathbf{-3.25}$	$\mathbf{-3.22}$	$\mathbf{-2.64}$	$\mathbf{-2.72}$
	(-4.21)	(-6.12)	(-5.63)	(-6.52)
$Vr_{i,t-1}(R)$	0.60	0.20	0.94	0.79
	(0.55)	(0.18)	(1.60)	(1.07)
$d^{tolerant}_{i,t-1}$	1.76	**5.51**	2.00	**7.28**
	(0.59)	(2.10)	(1.19)	(3.60)
$d^{\pi>\mathcal{T}_i}_{i,t-1}$	1.04		-4.28	
	(0.18)		(-1.29)	
$d^{tolerant,\mathcal{T}_i}_{i,t-1}$	4.65		**16.44**	
	(0.80)		(4.60)	
$d^{tolerant}_{i,t-1} \times \tilde{\pi}^{GAP}_{i,t-1}$		**3.90**		**1.66**
		(3.44)		(2.00)

Notes: see Table 4.

Table 6: Robustness analysis: Alternative specifications of the SUR model in (8) and (9)

Sample:	With RV_{t-1}	With $indep_{i,t-1}$	Before 2008	with $\bar{\mathbf{x}}_{t-1}$, $\lambda_t = 0$	$d_{i,t-1}^{tolerant} = 1$ f. Eurozone
$\pi_{i,t-1}$	−0.81	−1.07	−0.58	−1.04	−0.34
	(−1.19)	(−1.56)	(−0.89)	(−1.52)	(−0.50)
$y_{i,t-1}$	−0.01	0.01	−0.06	0.04	−0.04
	(−0.05)	(0.03)	(−0.37)	(0.26)	(−0.22)
$SD_{i,t-1}(\pi)$	**21.51**	**21.39**	**19.38**	**22.40**	**21.80**
	(9.21)	(9.25)	(8.45)	(9.84)	(9.39)
$SD_{i,t-1}(y)$	0.24	0.19	0.27	0.19	0.23
	(1.43)	(1.15)	(1.61)	(1.19)	(1.37)
RV_{t-1}	**16.71**				
	(6.14)				
$indep_{i,t-1}$		**−8.67**			
		(−3.03)			
$\tilde{R}^{+}_{i,t-1}$	−0.01	−0.08	0.59	−0.65	0.19
	(−0.02)	(−0.13)	(1.02)	(−1.06)	(0.32)
$\tilde{R}^{-}_{i,t-1}$	**−2.13**	**−2.20**	**−2.31**	**−2.44**	**−2.00**
	(−4.44)	(−4.61)	(−5.19)	(−4.99)	(−4.17)
$Vr_{i,t-1}(R)$	−0.09	−0.05	−0.35	0.42	−0.13
	(−0.15)	(−0.08)	(−0.67)	(0.71)	(−0.22)
$d_{i,t-1}^{tolerant}$	1.38	3.19	2.59	−0.46	−2.43
	(0.74)	(1.57)	(1.47)	(−0.25)	(−1.34)
$d_{i,t-1}^{\pi>\mathcal{T}_i}$	−1.80	0.51	−1.95	−2.47	−4.65
	(−0.51)	(0.14)	(−0.59)	(−0.74)	(−1.29)
$d_{i,t-1}^{tolerant,\mathcal{T}_i}$	**18.16**	**15.47**	**21.06**	**18.41**	**21.40**
	(4.76)	(3.96)	(6.12)	(5.03)	(5.64)

Notes: The fifth column shows results obtained when cross sectional averages of macroeconomic determinants, $\bar{\mathbf{x}}_{t-1} = (1/N)\sum_{i=1}^{N} \mathbf{x}_{i,t-1}$, are included in addition to the explanatory variables listed in the Table to control for potential cross sectional dependence in the disturbances of (8) (Pesaran, 2006). 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.

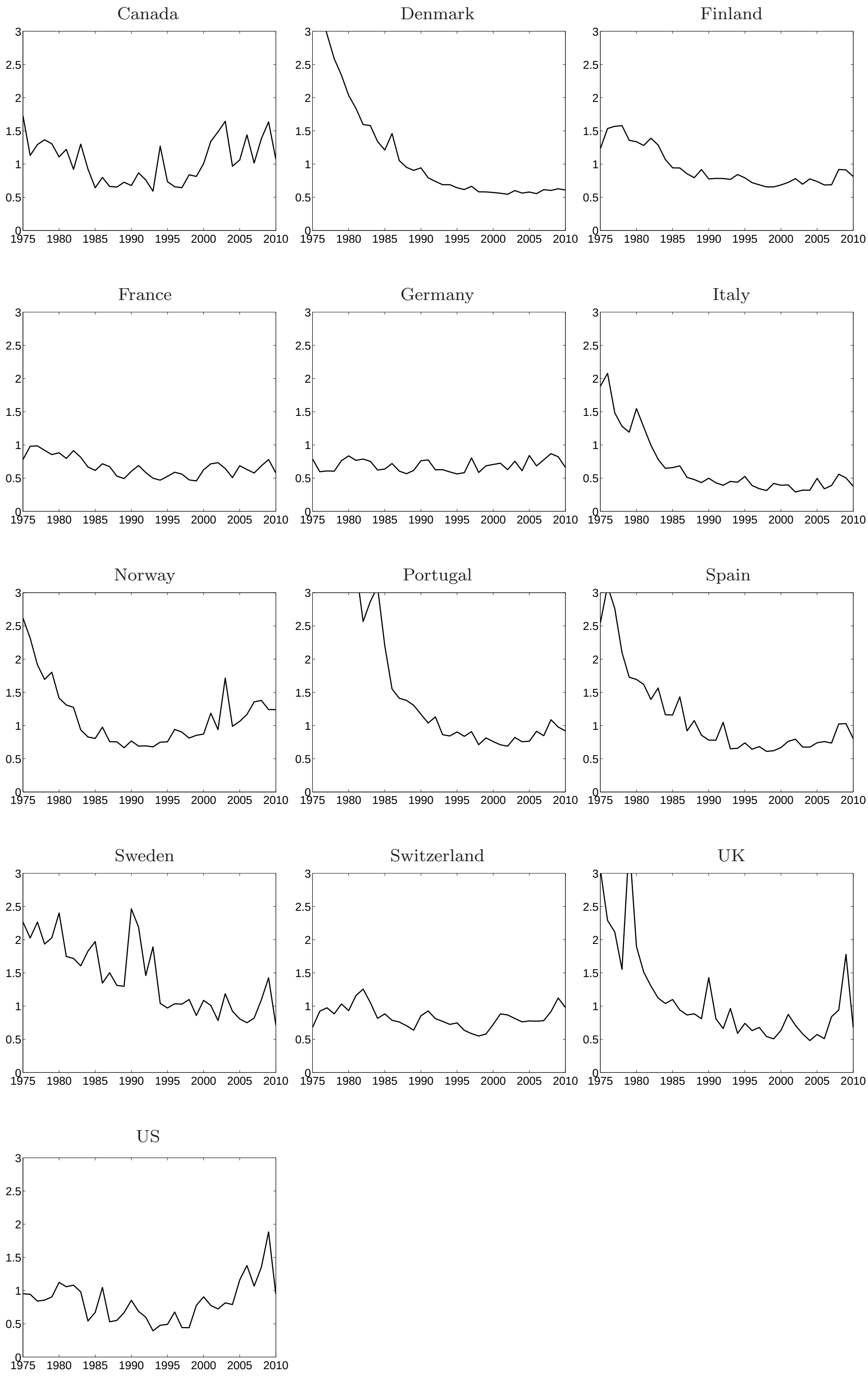

Figure 1: $IU_{i,t}$ from model (7) and (8) for 13 economies.

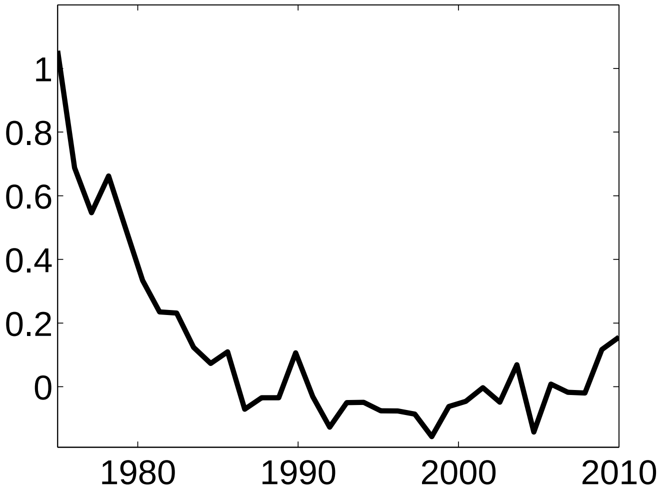

Figure 2: The estimated cross-sectional time trend in $IU_{i,t}$ as represented by $\hat{\lambda}_t$ in (9). Estimates $\hat{\lambda}_t$ are obtained from the model specification reported in the column I of Table 4.

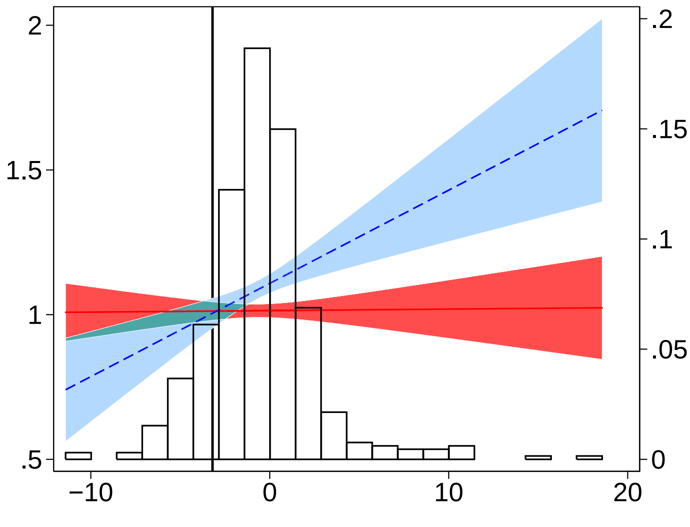

Figure 3: Predictions $\hat{IU}_{i,t}$ for $d^{tolerant}_{i,t-1} = 0$ (solid red line) and $d^{tolerant}_{i,t-1} = 1$ (dashed blue line), based on estimates from column X of Table 4. Shaded areas depict 95%-confidence intervals. The intersection of both predictions is indicated by the vertical line. Below, a histogram of $\tilde{\pi}^{GAP}_{i,t-1}$ is depicted. The magnitude of $\hat{IU}_{i,t}$ is measured on the left scale, the right scale corresponds to the values of the histogram for $\tilde{\pi}^{GAP}_{i,t-1}$.

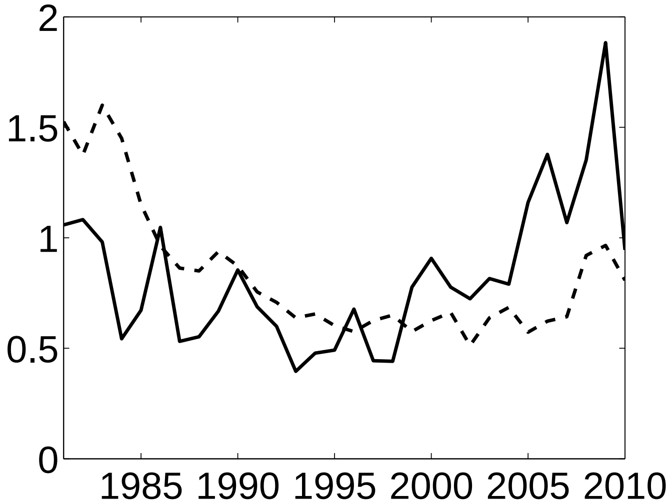

Figure 4: US inflation uncertainty as measured by $IU_{US,t}$ (Spline-GARCH, solid line) and $Dis_{US,t}$ (interquartile range of inflation forecasts from the Survey of Professional Forecasters, dashed line).