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Abstract

We investigate the international linkages of inflation uncertainty in the G7. In a first step, we document that inflation uncertainty in the G7 is intertwined. Moreover, the degree of synchronization has increased during the recent two decades. Second, based on a Factor-Structural Vector Autoregression (FSVAR) model, we provide evidence of a common international shock that drives national inflation uncertainty and which is closely related to oil and commodity price uncertainty. Third, we document that the size of shocks to inflation uncertainty has declined over time paralleling the process of inflation stabilization. Fourth, we analyze whether this decline can be explained by "good policy" or by "good luck". It appears that domestic shocks translate less extensively into the individual economies. We interpret this finding in favor of the "good policy" hypothesis. Finally, we document that the relative importance of international shocks has increased which explains the higher degree of synchronization in inflation uncertainty among the G7.

JEL-Code: C380, E310, E320.

Keywords: inflation uncertainty, Factor-Structural VAR, stochastic volatility.

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1 Introduction

It is well known that increased inflation uncertainty may lead to economic cost (see, for instance, Bernanke and Mishkin, 1997; Fischer and Modigliani, 1978). First, increased inflation uncertainty complicates the optimal decision regarding long-term savings and investment. Second, increased volatility may distort the signal of the price system. Third, nominal contracts involving, for instance, wages and financial assets become riskier.[1] Moreover, a strand of literature stresses that higher inflation uncertainty is typically associated with higher inflation (Ball, 1992; Cukierman and Meltzer, 1986; Friedman, 1977). As a consequence, inflation uncertainty increases the cost of high inflation and hampers the anchoring of low inflation expectations. Hence, understanding the evolution of inflation uncertainty is crucial if we want to maintain the benefits of low and stable inflation rates.[2]

Our study aims to provide additional insights into the international linkages of inflation uncertainty. The contribution of the present paper is twofold. First, we document the extent of co-movement of inflation uncertainty among the G7 and analyze the sources of international synchronization. Second, we investigate the origins of changes in the dynamics of national inflation uncertainty accounting for international factors and spillover effects from one country to the other. We tackle both questions with the help of a Factor-Structural Vector Autoregression (FSVAR) model which allows for a decomposition of the total variation of inflation uncertainty into the contributions of international shocks, own shocks, and spillover effects.

A number of studies focuses on common factors as a reason for business cycle synchronization (see, for instance, Kose et al., 2008; Stock and Watson, 2005). Likewise, incomplete exchange rate adjustment and exposure to global shocks such as, for example, oil-supply or commodity price shocks provide a basis for a common component in national inflation rates (see, for instance, Ciccarelli and Mojon, 2010; Mumtaz and Surico, 2012). Bataa et al. (2012b) analyze international linkages of inflation between major industrialized countries. They provide evidence of increased co-movement among the Euro area countries, as well as a rising correlation between the US, Canada and the Euro area aggregate. We extend this literature by analyzing the degree and the sources of synchronization of international inflation uncertainty.

[1]Recently, uncertainty shocks have also gained attention as drivers of business cycle fluctuations. A growing literature documents their potential effects on the real economy. See Bloom (2009), Alexopoulos and Cohen (2009), Fernandez-Villaverde et al. (2011), Bachmann and Elstner (2012), and Baker et al. (2012), among others.

[2]Consequently, a large number of empirical studies analyzes the effects of increased inflation uncertainty. Previous studies typically discuss its relation to inflation and output at the national level. See, for instance, Baillie et al. (1996), Grier and Perry (1998), Bhar and Hamori (2004), Fountas and Karanasos (2007), Fountas (2010), and Caporale et al. (2012).

We consider common shocks and spillover effects as possible explanations for synchronization of inflation uncertainty in the G7 and quantify the importance of each of these components for national inflation uncertainty.

Another strand of literature documents a decline in the volatility of inflation in the US since the mid-eighties (see Canova and Ferroni, 2012; Cogley et al., 2010; Stock and Watson, 2007, 2010, among others). Cecchetti et al. (2007) demonstrate that the volatility of trend inflation has decreased over time in the other G7 countries as well which constitutes an "Inflation Stabilization" process. Bataa et al. (2012a,b) analyze the nature and the timing of the changes in international inflation uncertainty by means of a statistical break test. In particular, for most G7 countries, they document a structural break in the volatility of inflation in the mid-eighties which is followed by a decline in inflation uncertainty.[3] In this paper, we investigate how the inflation uncertainty process has changed over time. We place an emphasis on the changes in the stability of inflation uncertainty. To shed light on the sources of these changes, we quantify the role of the size of shocks impinging on inflation uncertainty ("good or bad luck"). In addition, we assess to what extent changes in the structure of the economy and in the (monetary) policy stance have altered the propagation of these shocks ("good or bad policy").[4]

Our results can be summarized as follows. First, we find evidence of synchronization among inflation uncertainty in the G7 notably at business cycle frequencies. We show that the degree of synchronization has increased during the recent two decades. Second, we reveal a common shock that moves domestic inflation uncertainty in all G7 countries into the same direction. We find that this common shock is closely related to oil and commodity price uncertainty. By contrast, shocks originating in the US have an impact on a subset of countries only. Third, based on recursive estimations, we document that there has been a marked increase in the stability of inflation uncertainty, paralleling the "Inflation Stabilization" process. To the best of our knowledge this has not been documented elsewhere. Fourth, we document that the propagation mechanism of shocks to inflation uncertainty in the G7 has changed considerably over time. It appears that domestic shocks translate less extensively into the individual economies. We interpret this finding in favor of the "good policy" hypothesis. Finally, the relative importance of international shocks has increased over time which provides

[3] As documented in Bataa et al. (2012a), in Canada, the US, and (to a lesser extent) in the Euro area, the decline appears to be only temporary as the volatility of inflation shocks begins to rise in the late nineties again.

[4] A similar approach has been used by Ahmed et al. (2004), Stock and Watson (2005), Giannone et al. (2008), Justiniano and Primiceri (2008), Galí and Gambetti (2009), among others, to analyze the sources of the "Great Moderation" in the US.

an explanation for the higher degree of synchronization among the G7.

The paper is organized as follows. We introduce our measure of inflation uncertainty in section 2. In section 3, we examine the degree of synchronization of G7 inflation uncertainty and test for structural breaks in the inflation uncertainty process. The set-up of the FSVAR model is explained in section 4. The empirical results of the FSVAR estimation are presented in section 5. Section 6 summarizes and provides conclusions.

2 Measuring inflation uncertainty

Before turning to the main analysis we need to come up with a measure of unobserved inflation uncertainty. Ideally, uncertainty is derived from subjective probability density functions of decision makers. Such a measure relies on information about the subjective probability that future inflation will fall in a certain range. For the US, a number of studies use these types of uncertainty measures (see, for example, Zarnowitz and Lambros, 1987, Giordani and Söderlind, 2003, and Rich and Tracy, 2010). However, consistent data for a longer time span including all G7 countries are not available up to now.[5] This is why we opt for a model-based measure which has the advantage of being consistently available for a long history to analyze the international linkages among the G7.

In this study we use a stochastic volatility model which has recently been proposed to model uncertainty (see, for instance, Dovern et al., 2012; Fernandez-Villaverde et al., 2011). The stochastic volatility model – in contrast to a GARCH model – allows for a separate innovation impinging on volatility (see, for instance, Fernandez-Villaverde and Rubio-Ramirez, 2010). Moreover, Grimme et al. (2011) show that a measure based on stochastic volatility compares well with other (survey-based) measures of inflation uncertainty. We derive the measure from an unobserved component model with stochastic volatility (UC-SV). Notably, Stock and Watson (2007) show that the UC-SV model captures the salient features of inflation, and it is also very well suited as a forecast device. One reason for the comparatively good forecast performance is that it decomposes inflation into a stochastic trend and a transitory

[5]Subjective probability densities are provided for the US by the Survey of Professional Forecasters (SPF) maintained by the Federal Reserve Bank of Philadelphia, the ECB's Survey of Professional Forecasters which polls expectations about aggregate Euro area data, and the Survey of External Forecasters conducted by the Bank of England.

component. The UC-SV model is given by equations (1) to (5).

$$\pi_t = \bar{\pi}_t + \eta_t \qquad \eta_t \sim N(0, \sigma^2_{\eta,t}) \tag{1}$$

$$\bar{\pi}_{t+1} = \bar{\pi}_t + \epsilon_t \qquad \epsilon_t \sim N(0, \sigma^2_{\epsilon,t}) \tag{2}$$

$$\log \sigma^2_{\eta,t+1} = \log \sigma^2_{\eta,t} + \nu_{1,t} \tag{3}$$

$$\log \sigma^2_{\epsilon,t+1} = \log \sigma^2_{\epsilon,t} + \nu_{2,t} \tag{4}$$

$$\begin{pmatrix}\nu_{1,t} & \nu_{2,t}\end{pmatrix}' = N(0, \gamma I_2) \tag{5}$$

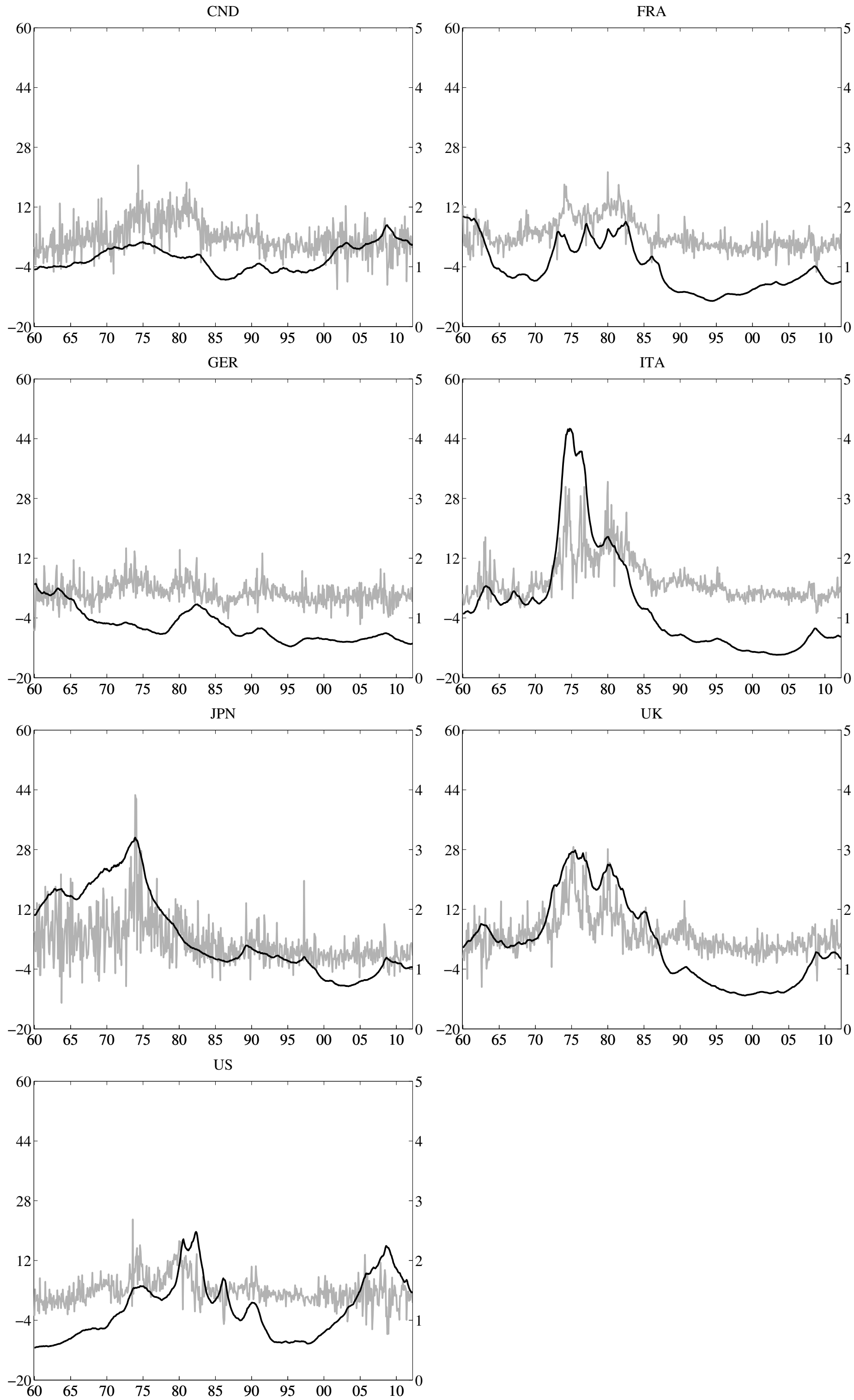
In this state-space model the trend $\bar{\pi}_t$ is modeled as a random walk which is driven by a level shock ϵ_t. The innovation process η_t captures the transitory part. The setting incorporates second moment shocks $\nu_{1,t}$ and $\nu_{2,t}$ which inflate the volatility of the process. The model is estimated with the Gibbs sampler.[6] An increase in the standard deviation of the permanent shock reflects that trend inflation is subject to larger changes which translate into larger forecast errors. Hence, $\sigma_{\epsilon,t}$ may be interpret as long-term inflation uncertainty. We believe that it is reasonable to assume that policy makers care more about the uncertainty associated with long-term inflation. Therefore, we follow Cecchetti et al. (2007) and focus on $\sigma_{\epsilon,t}$.

Our sample comprises the G7 (Canada, France, Germany, Italy, Japan, the United Kingdom, and the US) over the period 1960:M1-2012:M4. We measure inflation as the annualized monthly percent change in the Consumer Price Index (CPI) given by $1200 \times \log(CPI_t/CPI_{t-1})$. The inflation series are obtained from the OECD database and are seasonally adjusted. Finally, outliers in the data have been removed most of which are attributable to announced changes in the value-added tax rate.[7]

Figure 1 shows the uncertainty measures together with actual inflation. A similar pattern emerges for the G7. In light of the high inflation rates observed in the seventies, we measure a steady increase in inflation uncertainty during this time. The upswing is followed by a marked reduction in volatility of inflation rates in the mid-eighties which constitutes the process of "Inflation Stabilization" (Cecchetti et al., 2007). However, during the last decade, uncertainty has risen somewhat in the majority of the G7 economies. In particular, most uncertainty measures peaked again during the Global Financial Crisis (see also Clark, 2009, and Dovern et al., 2012 on this point).

[6] Estimation is based on the replication files of Stock and Watson (2007) which are available from Mark W. Watson's website: `http://www.princeton.edu/~mwatson/publi.html`. The model has only one scalar parameter γ which determines the smoothness of the stochastic volatility. Stock and Watson (2007) calibrate this parameter to $\gamma = 0.20$ for quarterly inflation rates. Since we have monthly data which usually carries more noise, we calibrate $\gamma = 0.2/3 = 0.07$.

[7] See appendix A for a detailed description of outlier adjustment.



Note: The grey line represents actual inflation (left-side axis), the dark line represents the long-term stochastic volatility measure of inflation uncertainty (right-side axis).

Figure 1: Inflation and long-term inflation uncertainty

3 Synchronization of inflation uncertainty in the G7

The first contribution of our study is to assess the degree of synchronization of inflation uncertainty among the G7. As proposed by Croux et al. (2001), we calculate the *dynamic correlation* between each country pair which shows the degree of synchronization at a given frequency. In the bivariate case, dynamic correlation between two variables x and y is defined as

$$\rho_{xy}(\lambda) = \frac{C_{xy}(\lambda)}{\sqrt{S_x(\lambda) S_x(\lambda)}}, \tag{6}$$

where $S_x(\lambda)$ and $S_y(\lambda)$ are the spectral density functions of x and y, $-\pi \leq \lambda < \pi$ is the frequency, and $C_{xy}(\lambda)$ is the cospectrum. The frequency λ is inversely related to the number of periods per cycle, $p = \frac{2\pi}{\lambda}$. Given our monthly data, a frequency of $\frac{\pi}{4}$, for example, corresponds to a cycle of 8 months.[8]

For a group of countries, co-movement can be summarized by the measure of cohesion which is defined as the (weighted) average of dynamic correlations among all possible country pairs:

$$coh_x(\lambda) = \frac{\sum_{i \neq j} w_i w_j \rho_{x_i x_j}(\lambda)}{\sum_{i \neq j} w_i w_j}, \tag{7}$$

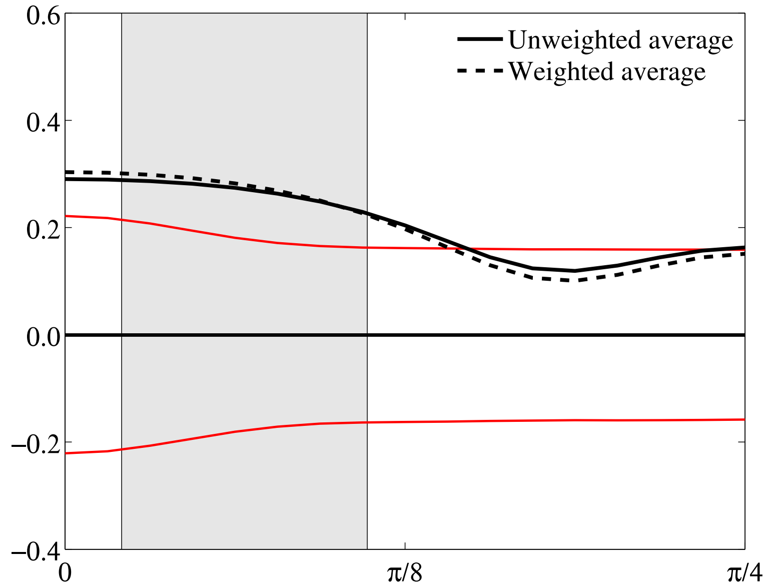
where x denotes a vector of variables, and w_i denotes the respective weight of country i. We consider two approaches. First, dynamic correlations are weighted equally with $w_i = 1$. Second, we use weights according to the country's share in the aggregate GDP of the G7.

Confidence bands for both, dynamic correlations and cohesion, are obtained from a bootstrapping procedure (see also Martin and Guarda, 2011). For each country pair, we calculate the dynamic correlation of two random normally distributed series of the same sample size and standard deviation as the original series. Based on 5,000 replications, we construct a confidence band at every frequency related to the null hypothesis that the two series are uncorrelated. Confidence bands for cohesion are given by the (weighted) average of the individual confidence bands.

The degree of synchronization as measured by cohesion is shown in figure 2. Cohesion is depicted at frequencies on the interval $[0,\pi/4]$, that is, from long-term cycles on the left-hand side up to the shortest cycle of 8 months on the right-hand side. The shaded area indicates the

[8] We depict the pairwise dynamic correlations in figure B in appendix B. It turns out that dynamic correlations at business cycle frequencies are positive and significant for most country pairs.

business cycle frequencies which typically cover 1.5 to 8 years. Generally, unweighted cohesion is significantly positive at business cycle frequencies since it remains above the confidence bands related to the null of no correlation. This finding suggests that business cycle frequencies contribute extensively to the co-movement of uncertainty measures across G7 countries. This result is also confirmed when we measure cohesion as a weighted average of G7 countries (dashed line in figure 2).



Note: The shaded area represents business cycle frequencies (8 to 1.5 years). Thin lines report 95% bootstrap confidence intervals. The weighted average is calculated according to the country shares in aggregate GDP of the G7 (based on values in US Dollars, constant prices and constant PPPs, OECD base year). The uncertainty measures were differenced beforehand. The Bartlett window size is set to 12.

Figure 2: Cohesion of inflation uncertainty among the G7

Given that our sample comprises more than fifty years of data, we may wonder whether there have been any major changes in the synchronization of inflation uncertainty. Moreover, one possible reason for the synchronization in the last fifty years may be that the countries experienced a common structural break. In the following, we assess whether there have been mean and/or variance breaks in the inflation uncertainty process. To this end, we conduct a standard sup-Wald test (Andrews, 1993) which also helps us to infer when a change occurs because it relies on the assumption of an unknown break date. For each country in our sample, we compute the Wald form of the Quandt likelihood ratio (QLR) statistic, maximized over the central 70% of the sample. The test for a mean break relies on an autoregressive model with twelve lags and the null hypothesis of constant autoregressive lag coefficients. To ensure non-negative values, we take the variables in logs. The test statistic for a break in the conditional variance is based on the null of a constant variance of the error term of the autoregressive model (see also Stock and Watson, 2002). The test allows for the possibility of two different break dates for the conditional mean and the conditional variance. The p-values corresponding to the QLR test statistics and the estimated break dates are reported in table 1.

	Conditional mean			Conditional variance		
	p-Value	Break date	67% confidence interval	p-Value	Break date	67% confidence interval
CND	0.05			0.41		
FRA	0.00	1993:12	1993:10 - 1994:02	0.00	1987:04	1986:12 - 1988:08
GER	0.00	1990:11	1990:09 - 1991:01	0.01	1987:05	1985:07 - 1990:10
ITA	0.01	1984:12	1984:10 - 1985:02	0.00	1977:03	1976:03 - 1979:12
JPN	0.00	1973:12	1973:10 - 1974:02	0.02	1996:03	1990:02 - 1998:06
UK	0.02	1984:11	1984:09 - 1985:01	0.04	1987:01	1983:12 - 1991:07
US	0.00	1973:03	1973:01 - 1973:05	0.04	1979:06	1969:11 - 1981:10

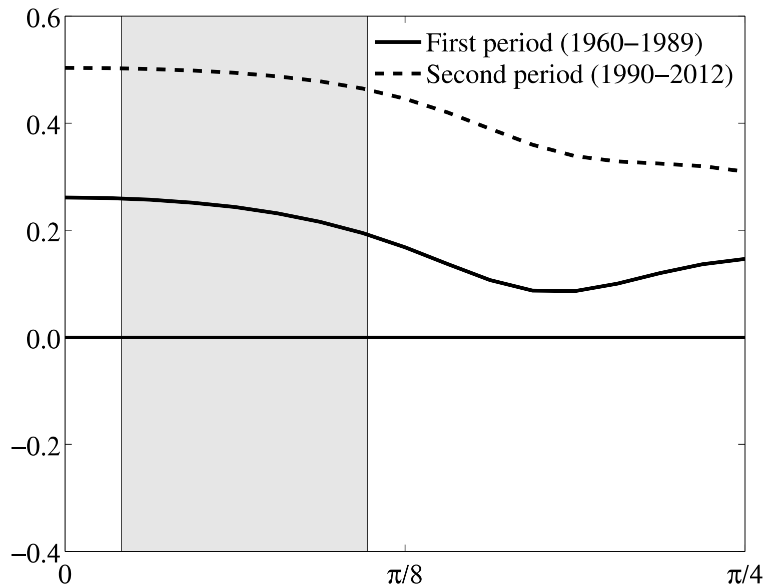
Note: Estimation based on AR(12) models for $\log \sigma_\epsilon$. The QLR test statistic on the "Conditional mean" refers to the null of no break in the AR lag coefficients. The test statistic on the "Conditional variance" refers to the null of no break in the variance of the AR error term. The break date and its confidence interval are estimated by OLS according to Stock and Watson (2005). Results are displayed only if the QLR test statistic is significant at least at the 5% level.

Table 1: Break tests for inflation uncertainty

We find evidence of breaks in the conditional mean as well as in the conditional variance for all countries. One exception is Canada where the null of no change in the mean and the null of no change in the AR innovations cannot be rejected at the 5% level. Concerning the conditional mean, the estimated break date is quite dispersed. Some countries experience a break during the first half of the seventies (Japan and the US), others in the mid-eighties (Italy and UK), and some in the beginning of the nineties (France and Germany). Hence, we cannot uniquely identify a (common) break date in the mean of inflation uncertainty. Notably, with the exception of France, the break occurs before the early nineties. Turning to breaks in the conditional variance, there seems to be some clustering for subgroups of countries. While variance shifts are detected in the late seventies in Italy and the US, the break in France, Germany, and the UK appears to occur in the late eighties. Again, most of the countries experience a break in the variance before the early nineties. Japan is somewhat outstanding. Here, a break in the conditional variance is indicated during the mid-nineties. Taken together, there have been marked changes in the dynamics of inflation uncertainty. However, it is difficult to identify a common break taking place in all countries synchronously. That is, the sources of the discrete breaks in the inflation uncertainty process appear to be country-specific and, hence, are not well suited as an explanation for the observed synchronization.

While we find evidence that inflation uncertainty in the G7 is intertwined in the full sample, an open question is whether synchronization of inflation uncertainty has changed over time.

Given the above break dates, it appears reasonable to split the sample in 1990, that is roughly in the middle. Since most countries have experienced a break before 1990, we compare synchronization, again measured by cohesion, before and after the change in the dynamics of inflation uncertainty. Figure 3 depicts cohesion calculated for the period 1960-1989 and 1990-2012, respectively. It becomes evident that cohesion increases considerably, that is, inflation uncertainty co-moves more strongly during the second sub-sample.[9]



Note: The shaded area represents business cycle frequencies (8 to 1.5 years).

Figure 3: Cohesion of inflation uncertainty, 1960-1989 and 1990-2012

4 The Factor-Structural VAR (FSVAR) model

The results presented in the previous section raise the question why uncertainty is synchronized in the G7 economies. In general, there might be two possible causes: common (global) shocks to inflation uncertainty and spillover effects from one country to the other. To disentangle both channels, we rely on a Factor-Structural VAR (FSVAR) model of the following form (see, for instance, Stock and Watson, 2005):[10]

[9]To test if the difference between the two sub-samples is statistically significant, we consider changes in the bivariate correlations obtained for the bandpass-filtered version of inflation uncertainty. Table C in appendix C shows the difference in pairwise correlations between the sub-samples 1960-1989 and 1990-2012. Evidently, the majority of pairwise correlations has increased significantly.

[10]The FSVAR set-up is also used by Altonji and Ham (1990), Norrbin and Schlagenhauf (1996), and Clark and Shin (2000) to model regional spillovers. For an application to international spillovers, see also Carare and Mody (2010) and Lahiri and Isiklar (2010), amongst others.

$$
\begin{aligned}
Y_t &= A(L)Y_{t-1} + v_t & (8)\\
v_t &= \Lambda f_t + \epsilon_t & (9)\\
E(v_t v_t') &= \Sigma_v & (10)\\
E(f_t f_t') &= diag(\sigma_{f_1}, ..., \sigma_{f_k}) & (11)\\
E(\epsilon_t \epsilon_t') &= diag(\sigma_{\epsilon_1}, ..., \sigma_{\epsilon_7}) & (12)
\end{aligned}
$$

Here, Y_t is a 7×1 vector stacking the demeaned uncertainty measures of the G7. The common factors are captured by f_t, Λ is the $7 \times k$ matrix of factor loadings, and ϵ_t denotes the idiosyncratic shocks. By assumption, the idiosyncratic shocks are uncorrelated with the common factors. The FSVAR model is estimated with Maximum Likelihood using the EM algorithm.[11] We set the lag-length to 12 which should be enough to capture the dynamics of the data. In order to ensure non-negative values of uncertainty, we take the log of $\sigma_{\epsilon,t}$.

According to equation (9) the error term of the FSVAR model is decomposed into country-specific idiosyncratic shocks and common shocks. Hence, a country-specific shock originating, for instance, in the US may be distinguished from a global shock. The global shock is identified by the assumption that it impacts all countries immediately whereas idiosyncratic shocks have an impact on other countries only via the autoregressive dynamics of the FSVAR model. We emphasize that, by using monthly data, we are less restrictive than previous studies dealing with business cycle synchronization based on quarterly data (see, for instance, Carare and Mody, 2010; Stock and Watson, 2005). In our case, spillovers are assumed to occur after one month already which implies that we attribute less explanatory power to the global shock(s) than studies based on quarterly data. However, if there are several global factors, these need to be identified separately. A common approach is to impose zero restrictions on the entries of Λ, the matrix of factor loadings (see, for instance, Gorodnichenko, 2006; Stock and Watson, 2005). We define Λ as an upper triangle where the first factor loads onto all G7 countries, the second factor has a zero restriction on the country ordered last (US), and the third factor has zero impact on both last ordered countries (UK and US).

In the next step, we have to pin down the number of common factors k which is achieved by testing the overidentifying restrictions of the model. The null hypothesis states that the FSVAR model has k common factors and 7 idiosyncratic shocks whereas the alternative states

[11] Estimations are based on the replication files of Stock and Watson (2005) which are available at Mark W. Watson's website.

that there are no restrictions imposed on the covariance matrix of the reduced-form errors v_t. The results of the corresponding Likelihood Ratio (LR) test are presented in table 2. For the sample 1960-2012, the test supports one common factor as the null of $k = 1$ cannot be rejected at the 5% level. However, this is a borderline case since one factor can still be rejected at the 10% level. We discuss the results of a specification with two common factors in section 5.1 later in the text and in appendix D.

k	logL (10^4)	d.f.	LR Stat.	p-value
0	20.5003	–	–	–
1	20.4897	14	21.30	0.09
2	20.4966	8	7.38	0.50
3	20.4991	3	2.48	0.48

Note: H_0: The reduced-form error covariance matrix has a k-factor structure. H_1: Unrestricted reduced-form error covariance. The number of overidentifying restrictions (d.f.) is given by $n(n+1)/2 - (nk - \sum_{j=0}^{k-1} j + n)$ where n is the number of equations in the FSVAR model.

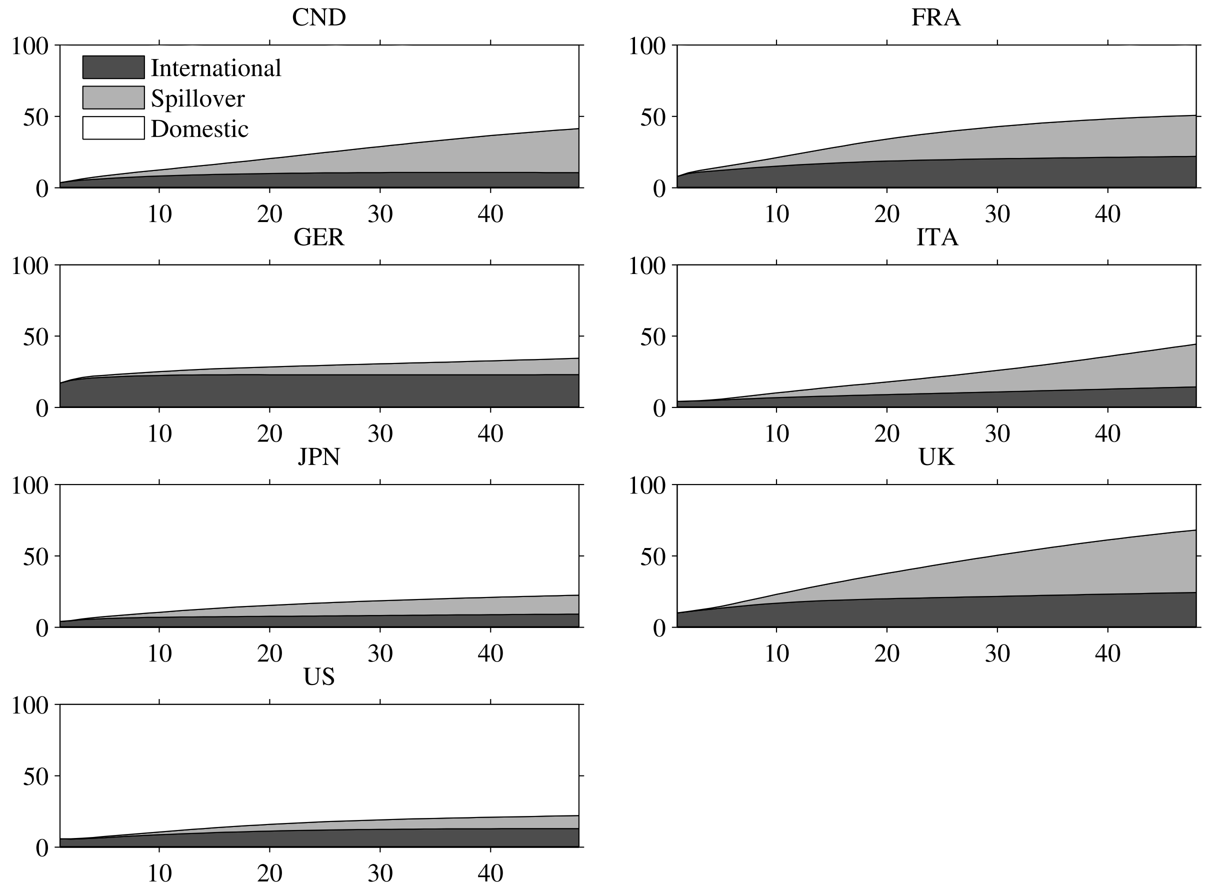
Table 2: Testing for the number of common factors in the FSVAR model

5 Empirical results

This section presents the empirical findings of the FSVAR estimation. We first analyze the importance of the three different types of shocks in each country. Second we investigate the impulse responses of the individual countries to the common shock as well as to a US shock. Third, we provide an economic interpretation of the common shock. Finally, we assess to what extent changes in the shock size or in the structure of the economy have altered the propagation of these shocks ("good luck" or "good policy").

5.1 How important are international shocks to inflation uncertainty?

In the following we assess the importance of the respective shocks impinging on inflation uncertainty with a particular focus on the international dimension. Since ϵ_t and f_t in the FSVAR model are uncorrelated by assumption, the forecast error variance for each country can be decomposed into international shocks, own shocks and spillovers received from other countries. Based on the FSVAR estimation, figure 4 displays the contribution of the three different types of shocks to the forecast error variance of inflation uncertainty at forecast horizons up to 48 months.



Note: The areas refer to the variance share of international shocks, spillovers, and domestic shocks, respectively. The forecast horizon in months is plotted on the abscissae. The results are based on an FSVAR model with one common factor and 12 lags.

Figure 4: Variance decomposition

The lower areas of figure 4 display the proportion of the international shock. The common factor has a noticeable impact on the Euro area countries and the UK. Table D in the appendix provides results for selected forecast horizons. At the two-year-horizon, the international factor captures up to 23% of the variance in the countries mentioned before. In contrast, for the North American countries and Japan, the common factor remains at about 10% across horizons. For Japan, the contribution of the international factor is the smallest of all countries.

The middle areas of figure 4 represent the proportion of spillovers. As noted before, spillovers do not contribute to the forecast error variance at the one-month horizon by assumption. Generally, the proportion of spillovers increases with the forecast horizon. At the one-year-horizon, the contribution of spillovers ranges between 3% and 8%.[12] In contrast, at the

[12]Note that this compares roughly with the contribution of monetary policy shocks to inflation documented

two-year-horizon, spillover-related shocks explain a comparably large part of the variance of inflation uncertainty, notably in Canada, France, Italy, and UK where the contribution ranges between 29% and 44%. In the US and Japan, the proportion of spillovers is somewhat smaller than in the other countries but comparable to the proportion of the international shock. The turbulent economic situation of Japan during the nineties and early 2000s associated with the asset price bubble seems to be reflected by a higher importance of domestic shocks. Likewise, Germany receives little spillover from abroad. This result is in line with the fact that the German Bundesbank has given very high priority to price stability during the entire sample period.

For most countries, the largest variance share is captured by domestic shocks, as reflected by the upper areas in figure 4. Generally, the proportion of domestic shocks is declining with the forecast horizon in favor of the international component (that is, in favor of common shocks or spillovers). At higher forecast horizons, the largest proportion of domestic shocks can be found in Japan and the US. Overall, the results of the variance decomposition suggest that domestic shocks play a major role for the variability of inflation uncertainty.

Given that one common factor may be rejected in favor of two common factors at the 10% level, we might wonder whether our results change when we introduce a second international shock. In table D in the appendix, we provide the results of the forecast error variance decomposition of an FSVAR model with two common factors. For the majority of countries, it turns out that the fraction of variance explained by the international component is practically unaffected if a second international shock is added to the model. Moreover, the fraction of own shocks hardly declines. Finally, the proportion explained by spillovers falls only marginally. A difference occurs, however, in the case of the US. Here, we observe an increase in the explained variance share; that is, the second shock adds explanatory power mainly for US inflation uncertainty.[13] Therefore, we believe that it is reasonable to proceed the analysis with one international shock.

5.2 Impulse response analysis

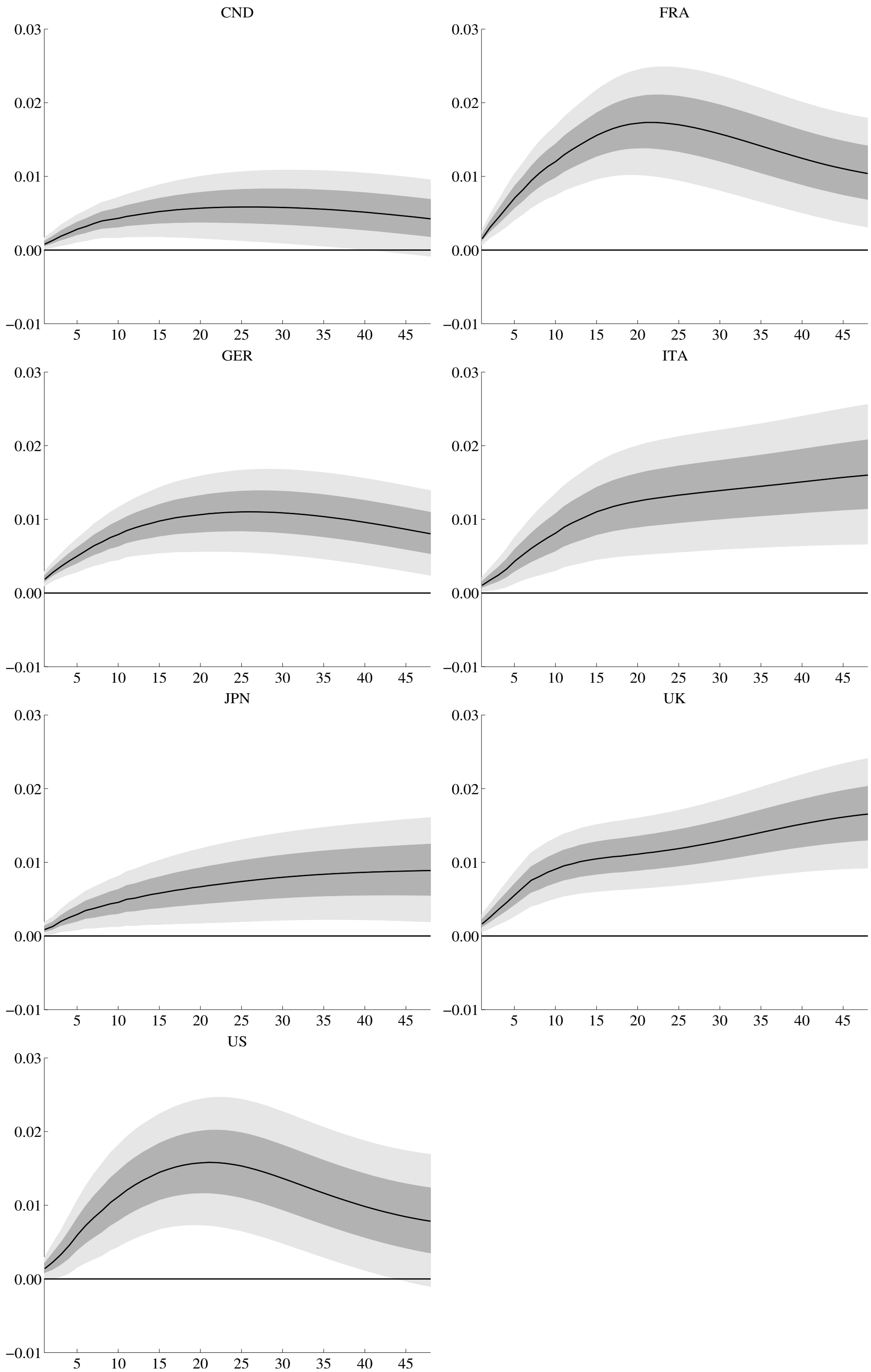
To see whether the common factor f_t qualifies for an international driver of inflation uncertainty, we calculate the response to a surprise increase in f_t. Figure 5 displays the impulse

in the literature (see, for instance, Bernanke et al., 2005; Christiano et al., 2005).

[13] The contribution of the two international shocks is also given separately in table D. The first international shock appears to be a shock that impinges on the US and – to a lesser extent – on Canada whereas the second international shock mainly affects the remaining countries. Note that this distinction is the result of the identification strategy concerning the factor loadings. The assumption of a recursive structure entails that the second common shock does not affect the US contemporaneously.

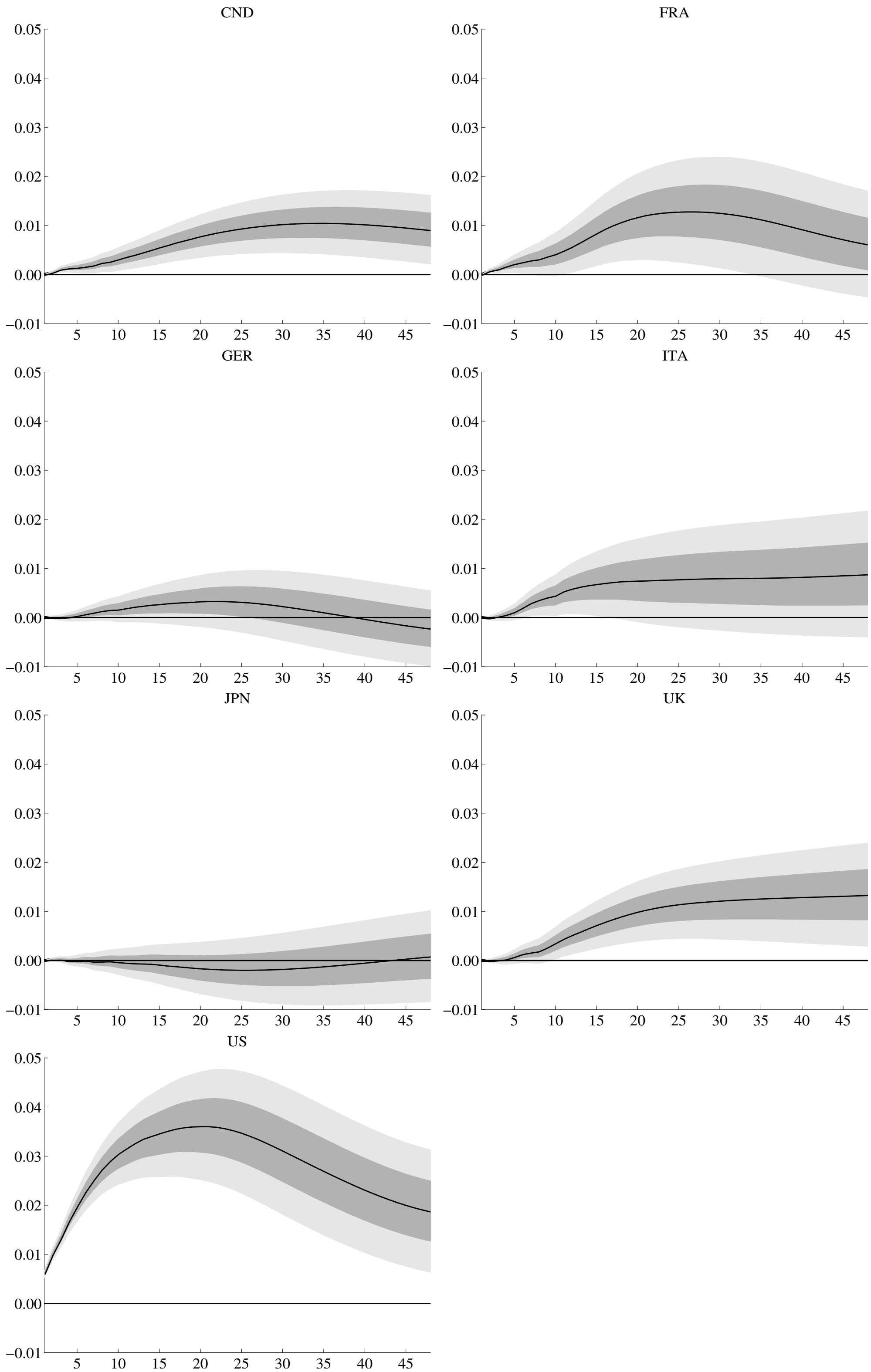
response functions of the individual countries following a one-standard deviation shock to the common factor. It turns out that a surprise innovation in the international factor shifts inflation uncertainty upwards in all countries. The impulse response follows a hump-shaped pattern, with a strong reaction in France, Italy, the UK and the US and a less pronounced increase in Canada, Germany and Japan. As the common shock uniformly drives uncertainty in the G7 economies into the same direction, f_t may be interpreted as an international shock to inflation uncertainty. Even more important, it provides a possible explanation for the synchronization which we document for these countries in section 3.

Also of interest is whether the international shock may be distinguished from a shock originating in the US. A sizeable impact of US shocks on the other G7 countries is, for example, documented by Bataa et al. (2012a). In the following, we analyze the effect of a surprise innovation in US inflation uncertainty. The impulse responses to a US shock are shown in figure 6. Note again that spillovers have no impact at the one-month horizon by assumption. Consequently, the contemporaneous effect of a shock to US inflation uncertainty is zero. In contrast to the common factor, the response to a US shock is mixed. In Canada, France, and the UK, a surprise innovation in US inflation uncertainty generates a significant rise which is, however, smaller and less persistent than the response to the international shock. For Germany and Japan, the reaction to the US innovation is insignificant. Hence, there seem to be country-specific differences, probably pertaining to the monetary regime, that determine how much inflation uncertainty spillover is received from the US. Given the countries' rather mixed reactions following a US shock, we conclude that the international factor can be distinguished from a shock to US inflation uncertainty. Moreover, the US innovation is only partly able to explain the synchronization among the G7.



Note: The bold black line represents the response of inflation uncertainty in the respective country to a one-standard deviation shock in the common factor. The shaded areas represent the 68% and 95% bootstrap confidence bands.

Figure 5: Response of inflation uncertainty to a shock to the common factor



Note: The bold black line represents the response of inflation uncertainty in the respective country to a one-standard deviation US shock. The shaded areas represent the 68% and 95% bootstrap confidence bands.

Figure 6: Response of inflation uncertainty to a US shock

5.3 Interpretation of the international shock

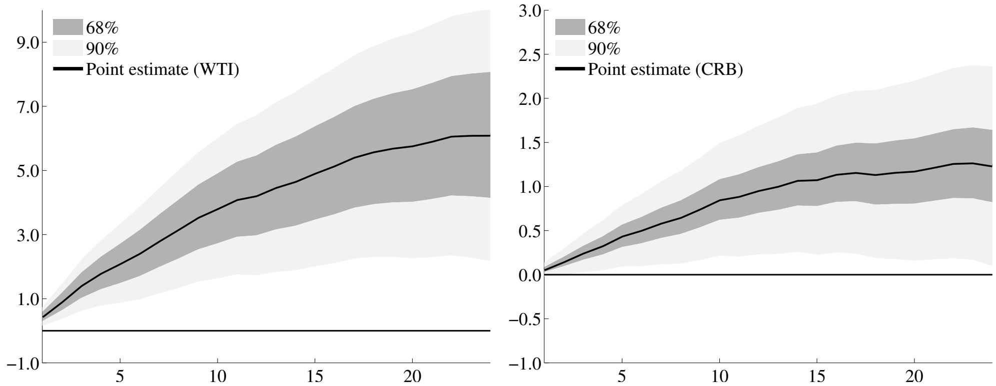
The following section aims to provide an economic interpretation of the international shock to inflation uncertainty. However, the FSVAR method does not allow for a direct interpretation of the common shock, and we are thus dependent on indirect evidence. A possible driver of international inflation uncertainty is the uncertainty associated with prices of goods which are traded all over the world at a common price, and which have a non-negligible share in the overall price index. Candidates are oil and commodity prices. Consequently, we would expect that the uncertainty related with those variables is foreshadowed by positive common shocks. We can infer whether f_t and any other measure of uncertainty are related by estimating the following regression:

$$unc_t^i = c_0 + \sum_{j=0}^{J} \phi_j^i f_{t-j} + \nu_t^i. \tag{13}$$

Here, unc_t^i represents a measure of uncertainty, and ν_t^i is the respective regression residual. Since f_t is an orthogonal white noise process by assumption, the coefficients ϕ_j^i are a measure of connectedness between unc_t^i and the common shock. Moreover, owed to the simplicity of the regression model, the cumulative coefficients can be interpreted as the impulse response of unc_t^i following a one percent increase in f_t (see, for instance, Kilian, 2009; Romer and Romer, 2010). We estimate the model with $J = 24$ lags.

To obtain a measure of oil price uncertainty, we use the UC-SV model introduced in section 2 and apply it to the monthly growth rate of the spot price for crude oil (WTI). The estimation sample runs from 1979M6 until 2012M4 since there is practically no monthly variation in WTI oil prices before that period. In addition, we also use the CRB/Reuters commodity price index and derive a measure of overall commodity price uncertainty in the same way. The CRB is more comprehensive than the WTI oil price since it measures the price of a basket of different commodities. Moreover, the CRB is available for the entire sample period (1960M1 until 2012M4). In order not to run into stationarity problems, unc_t^i is the log-change of the respective standard deviation associated with the long-term component of oil or commodity price inflation. In figure 7 we depict the dynamics of unc_t^i following an increase in the common shock f_t.

It appears that both oil and commodity price uncertainty are connected with the international shock to inflation uncertainty. Notably, increases in the international shock seem to



Note: The solid line represents the response of unc_t^i to a one percent increase in f_t. The 68% and 90% error bands are obtained by a block bootstrap using a block size of 12 and 20,000 replications. The left panel depicts the response of WTI oil price uncertainty while the right panel depicts the response of CRB commodity price uncertainty.

Figure 7: Response of oil and commodity price uncertainty to f_t

foreshadow increases in oil and commodity price uncertainty.[14] Taken together, our results provide evidence that f_t may be interpreted as a shock to international commodity price uncertainty which shows up as a common shock to inflation uncertainty in the G7.

5.4 Changes in the dynamics of international inflation uncertainty

Up to now, we have based the analysis on the full sample period. However, the break test in section 3 already indicated that there have been changes to the inflation uncertainty process during the last fifty years. In what follows, we investigate whether and how the dynamics of inflation uncertainty in the G7 have changed over time. To this end, we perform a recursive estimation of the FSVAR model with one common factor. A time-varying specification based on the whole sample is obtained by two-sided exponential weighting (Bataa et al., 2012b; Stock and Watson, 2005). That is, the regression at time t is calculated using weighted observations. The observation at time s receives an exponentially decreasing weight $\delta^{|t-s|}$ with a discount factor $\delta = 0.97$. Running $t = 1 \dots T$ over the whole sample, we obtain a time-varying estimate for the volatility of inflation uncertainty measured by the total forecast error variance. Note that the weighted estimation implies a smooth transition of the coefficients over time. It is thus not possible to retrace sudden breaks in the data. Nevertheless, the method provides us with an intuition whether there have been changes over the full sample

[14]In appendix E we also run the regression in equation (13) for a measure of financial market uncertainty used in Bloom (2009). We find no significant relation between the international factor and financial market uncertainty.

period.

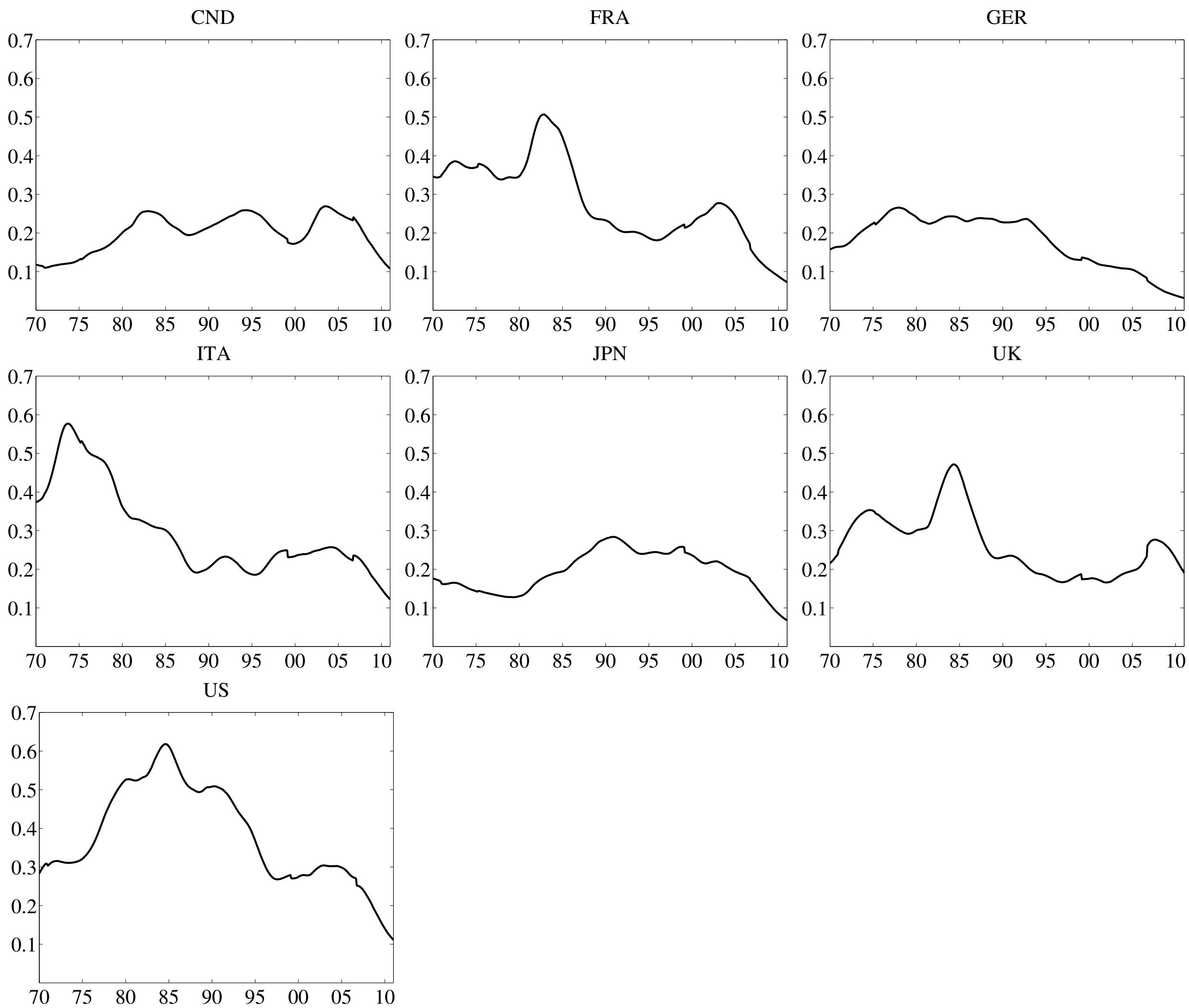
The time-varying forecast error variance of inflation uncertainty is depicted in figure 8 from 1970M1 onwards. In most countries, we observe a decline in volatility taking place in the mid-eighties to the early nineties. Except for Germany and Japan, a temporary increase in volatility can be observed in the beginning of the millennium. Towards the end of the sample, the variance has decreased to historically low levels in all countries. The timing of the observed changes suggests that the decline of the fluctuations of uncertainty has paralleled the process of world-wide "Inflation Stabilization" (see, for instance, Cecchetti et al., 2007; Mumtaz and Surico, 2012). That is, inflation uncertainty has not only come down to lower levels but is now also much more stable and, hence, easier to predict.

In general, two possible explanations are at hand for the documented increase in the stability of inflation uncertainty: either the size of the shocks has decreased ("good luck") or structural changes in the economy have dampened the transmission of the shocks ("good policy"). In the following, we decompose the decline of the total forecast error variance into the effect of the shock size and the propagation of the different types of shocks. Given the break dates obtained in section 3, we split the sample roughly in half and consider the periods 1960-1989 and 1990-2012. This allows us to compare the forecast error variance decompositions for the two different sub-samples. Moreover, we are able to distinguish between changes in the shock size and changes in the impulse response. Let V_p denote the variance of the forecast error, where $p = 1, 2$ refers to the first and second sub-sample, respectively. For notational simplicity, we suppress the dependence on the forecast horizon and the country. Since the FSVAR model incorporates eight sources of variation (one international shock, one domestic shock, and six different spillover terms emerging from the idiosyncratic shocks), the total variance can be written as $V_p = V_{p,1} + ... + V_{p,8}$ with $V_{p,j}$ denoting the contribution of the jth shock in sub-sample p. Consequently, the difference between the first- and the second-period variance can be expressed as $V_2 - V_1 = (V_{2,1} - V_{1,1}) + ... + (V_{2,8} - V_{1,8})$.

The variance of the forecast error consists of two parts: $V_{p,j} = a_{pj}\sigma^2_{pj}$, where a_{pj} is given by the cumulated squared impulse responses to a standardized (unit) shock j. σ^2_{pj} denotes the variance of shock j in sub-sample p. For each shock j, the change in the contribution to the total variance can be expressed as

$$V_{2j} - V_{1j} = \left(\frac{a_{1j} + a_{2j}}{2}\right)(\sigma^2_{2j} - \sigma^2_{1j}) + \left(\frac{\sigma^2_{1j} + \sigma^2_{2j}}{2}\right)(a_{2j} - a_{1j}). \tag{14}$$

The first term of the right-hand side in equation (14) refers to the contribution from the



Note: The bold line refers to the 12-months-ahead total forecast error variance of inflation uncertainty. Results are derived from a recursively weighted estimation of an FSVAR model with one common factor and 12 lags.

Figure 8: Time-varying volatility of inflation uncertainty

change in the shock size whereas the second term refers to the contribution from the change in the impulse response function. Table 3 reports the decomposition of changes in the 12-month-ahead forecast error variance of inflation uncertainty. Focusing on the first panel, the total variance has decreased significantly in most countries during the second period, as shown in column (3). That is, inflation uncertainty has become more stable and its predictability has increased. In Canada and Japan, the forecast error variance has slightly increased; however, this increase is statistically insignificant. Overall, the decline in volatility inferred from visual inspection of the recursive FSVAR estimation above can be recovered when we split the sample.

	Total variances			**Contribution of change in shock size**				**Contribution of change in impulse responses**			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	1960-1989	1990-2012	Change	Common	Spillover	Domestic	Total	Common	Spillover	Domestic	Total
CND	2.74^{***}	3.26^{***}	0.53	0.05	-0.14^{*}	0.33	0.24	0.08	0.05	0.16	0.29
	(0.37)	(0.45)	(0.58)	(0.13)	(0.08)	(0.30)	(0.35)	(0.42)	(0.22)	(0.39)	(0.66)
FRA	24.75^{***}	4.36^{***}	-20.39^{***}	0.40	-0.41^{*}	-7.72^{***}	-7.73^{***}	-5.20^{***}	-0.37	-7.09^{***}	-12.66^{***}
	(3.25)	(0.61)	(3.30)	(0.61)	(0.21)	(1.39)	(1.47)	(1.58)	(0.68)	(1.70)	(2.63)
GER	7.16^{***}	1.92^{***}	-5.24^{***}	0.15	-0.14^{*}	-0.77	-0.76	-2.06^{***}	-0.09	-2.33^{***}	-4.48^{***}
	(0.96)	(0.25)	(0.99)	(0.24)	(0.08)	(0.48)	(0.48)	(0.73)	(0.23)	(0.62)	(0.97)
ITA	18.98^{***}	5.80^{***}	-13.19^{***}	0.09	-0.05	-5.84^{***}	-5.80^{***}	-1.19	-0.79	-5.40^{***}	-7.39^{***}
	(2.56)	(0.78)	(2.67)	(0.29)	(0.15)	(1.23)	(1.31)	(1.07)	(0.56)	(1.76)	(2.29)
JPN	2.88^{***}	3.47^{***}	0.59	0.02	-0.26^{**}	0.58^{**}	0.34	-0.21	0.52^{*}	-0.06	0.25
	(0.38)	(0.47)	(0.60)	(0.07)	(0.11)	(0.27)	(0.31)	(0.26)	(0.31)	(0.45)	(0.63)
UK	8.72^{***}	3.89^{***}	-4.83^{***}	0.13	-0.29^{*}	-0.56	-0.72	-1.41	-0.90^{*}	-1.81^{**}	-4.11^{***}
	(1.11)	(0.58)	(1.26)	(0.30)	(0.18)	(0.74)	(0.71)	(0.92)	(0.53)	(0.78)	(1.37)
US	18.52^{***}	5.87^{***}	-12.66^{***}	0.31	-0.17	-8.53^{***}	-8.38^{***}	2.87^{**}	-0.91	-6.23^{***}	-4.27^{*}
	(2.52)	(0.82)	(2.66)	(0.42)	(0.18)	(1.85)	(1.71)	(1.15)	(0.61)	(1.72)	(2.49)

Note: The left panel shows the 12-months-ahead forecast error variance of inflation uncertainty for two sub-samples and the difference between the two sub-samples based on an FSVAR estimation with one common factor and 12 lags. The middle panel shows the contribution of the shock size to the changes in column (3). The right panel shows the contribution of changes in the impulse responses to the changes in column (3). The columns (7) and (11) add up to the total change shown in column (3). The values are multiplied by 100 and bootstrap standard errors are reported in parentheses.

Table 3: Decomposition of changes in the forecast error variance

The second panel of table 3 reports the contribution to the change in the forecast variance by *changes in the shock size*. The decline in the forecast variance is partly due to smaller shocks, as shown in column (7). This negative change in the shock size is statistically significant for France, Italy, and the US. The presence of smaller shocks suggests that "good luck" has contributed to the decline of the volatility of inflation uncertainty in these countries. Concerning the different shock types, mainly domestic shocks account for a decline in the shock size. In the majority of countries smaller spillovers also contributed to this downward trend in a significant way. By contrast, international shocks are larger in the second period, although this change is not statistically significant.

The third panel of table 3 displays the contribution of *changes in the impulse responses*, that is, changes in the way shocks translate into the domestic economy. As reported in column (11), the sensitivity towards shocks has generally decreased. In all countries except Canada and Japan, changes in the impulse responses significantly contributed to the overall decline in variance. It appears that "good policy" is responsible for most of the decline in the volatility of inflation uncertainty. The total contribution of the propagation mechanism can be further decomposed into the contribution of the transmission of the international shock, spillovers from other countries, and own shocks. Results regarding the international shock are rather mixed. In France and Germany, the effect of the international shock becomes weaker whereas the effect becomes significantly stronger in the US. The propagation mechanism of spillovers remains largely unchanged as the estimated difference is insignificant in most countries. Finally, most G7 countries became less sensitive to own shocks. Notably, the change in the propagation of own shocks accounts for the largest part of the decline reported in column (11).

The main message from table 3 is twofold. First, primarily the change in the propagation mechanism of shocks to inflation uncertainty has contributed to a "moderation" in inflation uncertainty. Second, the size of domestic shocks to inflation uncertainty has decreased in many countries while international shocks have become slightly larger. Taken together, the international shock has gained relative importance. Hence, increased synchronization is the result of own shocks losing importance relative to the international factor.

The above results raise the question what is so different in the second period that led to a stabilization of inflation uncertainty? One policy area that underwent major changes in the last two decades is the field of monetary policy. Researchers seem to agree that, in the developed countries, there is now a better understanding of how to implement monetary policy (see, for instance, Cecchetti et al., 2006; Summers, 2005). In particular in the US, monetary policy moved from an accommodative stance to an inflation-stabilizing policy (see, among others, Clarida et al., 2000). In a related empirical exercise, Ahmed et al. (2004) document that most of the decline in inflation volatility in the US may be explained by better monetary policy. Similarly, for the G7, Cecchetti et al. (2007) argue that the world-wide "Inflation Stabilization" is strongly linked to central banks acting more responsive to inflationary shocks. Accompanying the policy shift, there have been a number of institutional changes in monetary policy in the G7. Major legislative reforms that enhanced central bank independence were adopted in France, Italy, Japan, and the UK during the nineties. Measured by different indexes of political and economic autonomy, central bank independence has generally risen in G7 countries from the first half of our sample to the second half since 1990 (see Acemoglu et al. 2008 and Arnone et al. 2009). Institutional changes during the last two decades also comprise the announcement of an inflation target. Among the G7, Canada and the UK introduced an official inflation target in the early nineties whereas the EMU countries adopted the ECB's quantitative target of price stability "below, but close to, 2% over the medium term".[15] Among others, Mishkin and Schmidt-Hebbel (2007) document that countries following an inflation targeting strategy successfully improve their macroeconomic performance by providing a strong nominal anchor. They also stress that inflation targeting countries are less sensitive to international shocks by strengthening domestic monetary policy independence. Our results suggest that the above policy changes did not only reduce inflation uncertainty in the last two decades but also contributed to a stabilization of inflation uncertainty.

6 Concluding remarks

Our study aims to provide additional insights into the international linkages of inflation uncertainty. For this purpose, we use monthly CPI inflation rates in the G7 from 1960 onwards. Our results can be summarized as follows. First, we find evidence of synchronization among inflation uncertainty in the G7. We show that the degree of synchronization has increased during the recent two decades. Second, in an FSVAR framework, we reveal a common shock that moves national inflation uncertainty in all countries into the same direction. We find that this common shock is closely related to oil and commodity price uncertainty. By contrast, a pure US shock induces mixed responses in the G7. Third, based on recursive estimations, we show that the volatility of inflation uncertainty has decreased over time paralleling the process of "Inflation Stabilization". Fourth, we document that the propagation mechanism of shocks to inflation uncertainty in the G7 has changed considerably since 1990. The main channel for the decline of the volatility of inflation uncertainty seem to be domestic shocks that translate less extensively into the individual economies. This finding supports the hypothesis of "good policy". Finally, there appears to be a higher connectedness of inflation uncertainty among the G7 which is traceable to an increase in the relative importance of international shocks.

As stressed by Cecchetti et al. (2007), the main candidate for the "Inflation Stabilization" in the G7 are changes in the monetary regime. This also provides a candidate explanation for the

[15] Since the beginning of 2012, that is, towards the end of our sample, the US and Japan also communicate an inflation target.

observed "moderation" in inflation uncertainty. Although inflation uncertainty is currently rather stable, we should bear in mind that this appears to be the result of central banks which credibly fight inflation. Moreover, accepting higher inflation – as recently called for to deal with the problem of excessive debt – may bring about the additional cost of higher world wide inflation uncertainty via spillover effects.

Appendix

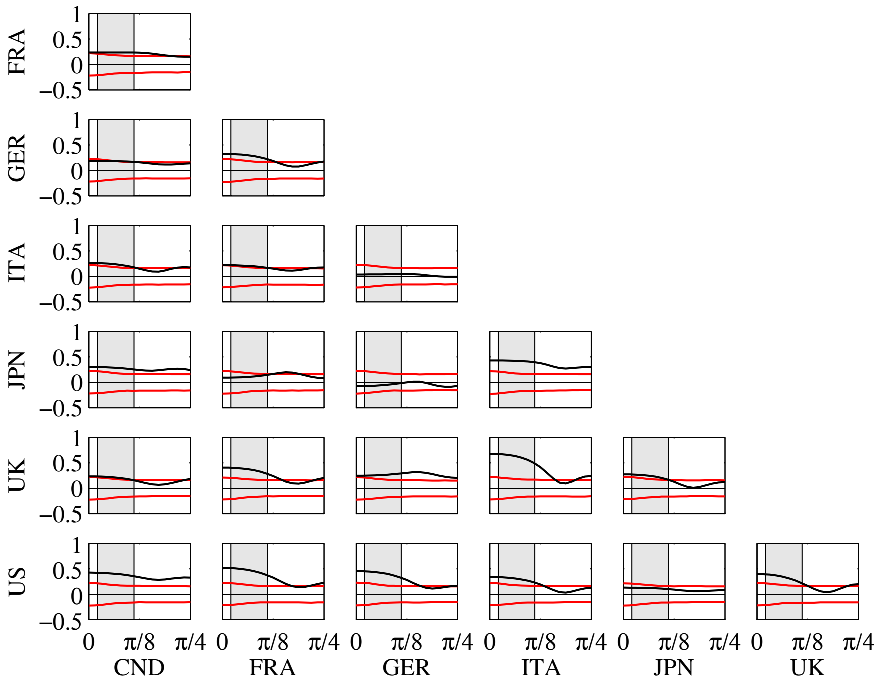
A Outlier adjustment

Before the analysis is conducted, we remove a number of outliers in the seasonally adjusted monthly inflation rates. Table A summarizes the adjustment of outliers. First, we identified outliers which are traceable to changes in the tax system; in most cases, we identify an increase in the value-added tax rate. The outliers in France in 1965M6 and 1965M7 are due to one exceptional observation in the level series of the CPI. Second, following Stock and Watson (2002), we refer to an outlier in the data if an observation deviates more than six times the interquartile range from the local mean. These outliers are marked with an asterisk in table A. All outliers are replaced with the mean of the six adjacent observations.

Canada		France		Germany	
1991M01	goods and services tax	1965M06	–	1991M10	German reunification
1994M01	arctic outbreak	1965M07	–	1993M01	VAT rate from 14% to 15%
1994M02	severe spending cuts				
UK		US			
1975M05*	–	2008M11*	–		
1979M07	VAT rate from 8% to 15%				
1991M04	VAT rate from 15% to 17.5%				

Table A: Adjustment of outliers in inflation

B Dynamic correlations among country pairs



Note: The shaded area represents business cycle frequencies (8 to 1.5 years). Thin lines report 95% bootstrap confidence intervals. The uncertainty measures were differenced beforehand. The Bartlett window size is set to 12.

Figure B: Dynamic correlation of inflation uncertainty in the G7 countries

C Testing for changes in correlations among country pairs

	Difference between 1990-2012 and 1960-1989					
	CND	**FRA**	**GER**	**ITA**	**JPN**	**UK**
FRA	0.41^{**}					
	(0.19)					
GER	0.36^{**}	0.32^{**}				
	(0.17)	(0.14)				
ITA	0.44^{*}	0.53^{**}	0.13			
	(0.24)	(0.22)	(0.23)			
JPN	0.11	0.16	0.29^{*}	0.01		
	(0.29)	(0.24)	(0.16)	(0.31)		
UK	0.62^{***}	0.39^{*}	-0.03	0.31^{**}	0.21	
	(0.15)	(0.22)	(0.13)	(0.16)	(0.21)	
US	0.48^{***}	0.35^{**}	0.05	0.47^{***}	0.25	0.63^{***}
	(0.15)	(0.17)	(0.20)	(0.17)	(0.19)	(0.19)

Note: The entries indicate the difference in correlation between the two sub-samples. Newey-West standard errors robust to heteroskedasticity and autocorrelation up to 12 lags are reported in parentheses. Uncertainty measures were detrended by means of a bandpass filter which extracts business cycle frequencies (1.5 to 8 years).

Table C: Differences in pairwise correlations of inflation uncertainty

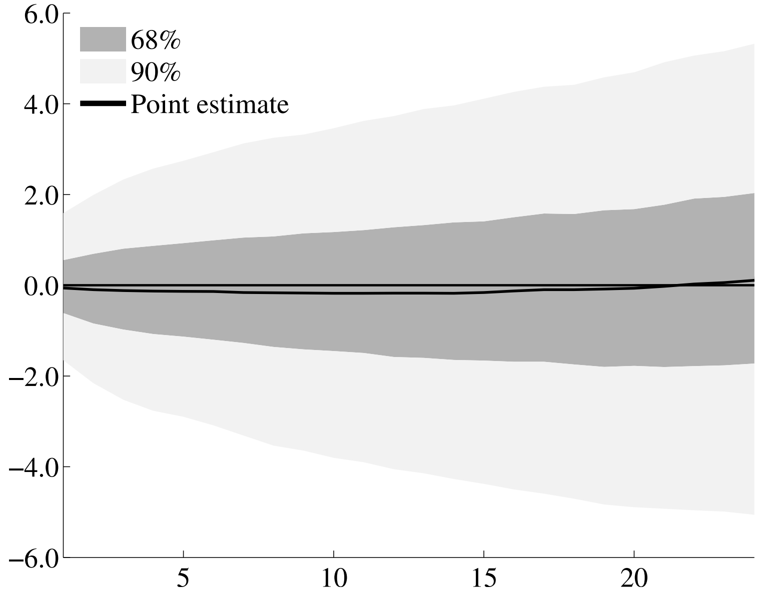
D Number of common factors in the FSVAR model

		One common factor				Two common factors					
		FE	Fraction of FEV due to:			FE	Fraction of FEV due to:				
	Horizon	STD	Int.	Spillovers	Own	STD	Int.	Int. 1	Int. 2	Spillovers	Own
CND	1	0.00	0.04	0.00	0.96	0.00	0.04	0.03	0.01	0.00	0.96
	12	0.25	0.09	0.05	0.86	0.25	0.12	0.09	0.03	0.04	0.84
	24	0.86	0.10	0.13	0.76	0.87	0.16	0.13	0.03	0.10	0.74
	48	2.87	0.11	0.31	0.59	2.91	0.21	0.19	0.03	0.22	0.56
FRA	1	0.01	0.08	0.00	0.92	0.01	0.07	0.03	0.04	0.00	0.93
	12	0.45	0.16	0.08	0.76	0.46	0.16	0.07	0.09	0.07	0.77
	24	1.73	0.19	0.19	0.62	1.73	0.21	0.09	0.11	0.17	0.62
	48	5.44	0.22	0.29	0.49	5.44	0.25	0.13	0.13	0.25	0.50
GER	1	0.00	0.17	0.00	0.83	0.00	0.15	0.03	0.13	0.00	0.85
	12	0.28	0.23	0.03	0.74	0.28	0.21	0.05	0.17	0.03	0.75
	24	1.06	0.23	0.06	0.71	1.06	0.22	0.05	0.17	0.06	0.72
	48	3.59	0.23	0.11	0.66	3.59	0.22	0.05	0.17	0.11	0.67
ITA	1	0.01	0.04	0.00	0.96	0.01	0.05	0.00	0.05	0.00	0.95
	12	0.43	0.07	0.05	0.88	0.43	0.09	0.01	0.08	0.04	0.87
	24	1.68	0.10	0.11	0.79	1.68	0.12	0.02	0.10	0.10	0.78
	48	5.39	0.14	0.30	0.56	5.39	0.18	0.04	0.14	0.28	0.55
JPN	1	0.00	0.04	0.00	0.96	0.00	0.09	0.00	0.09	0.00	0.91
	12	0.28	0.07	0.04	0.88	0.28	0.12	0.00	0.12	0.05	0.84
	24	1.08	0.08	0.09	0.83	1.08	0.12	0.00	0.12	0.09	0.79
	48	3.84	0.09	0.13	0.78	3.84	0.13	0.00	0.13	0.13	0.74
UK	1	0.01	0.10	0.00	0.90	0.01	0.13	0.00	0.13	0.00	0.87
	12	0.34	0.18	0.08	0.74	0.34	0.21	0.02	0.19	0.08	0.71
	24	1.21	0.21	0.22	0.57	1.21	0.25	0.05	0.20	0.20	0.55
	48	4.08	0.24	0.44	0.32	4.05	0.30	0.10	0.20	0.38	0.31
US	1	0.01	0.06	0.00	0.94	0.01	0.44	0.44	0.00	0.00	0.56
	12	0.52	0.09	0.03	0.88	0.52	0.44	0.44	0.01	0.02	0.53
	24	2.01	0.12	0.06	0.83	1.99	0.44	0.42	0.02	0.06	0.50
	48	6.28	0.13	0.09	0.78	6.18	0.43	0.40	0.03	0.09	0.48

Note: The table reports the standard deviation (STD) and the variance decomposition of inflation uncertainty forecast errors at the 1-, 12-, 24-, and 48-months horizon. Fractions are given as percentage of total forecast error variance (FEV). Estimation based on an FSVAR model with 12 lags.

Table D: Variance decomposition into international shocks, spillovers, and domestic shocks

E Financial market uncertainty and the common shock f_t



Note: The solid line represents the response of financial market uncertainty to a one percent increase in f_t. Financial market uncertainty is the log of the uncertainty measure in Bloom (2009) who uses the VXO and the VIX from the Chicago Board Options Exchange to construct a long time series of financial market uncertainty beginning in 1962M8. For the estimations in the present paper we have updated Bloom's series until 2012M4. The 68% and 90% error bands are obtained by a block bootstrap using a block size of 12 and 20,000 replications.

Figure E: Response of financial market uncertainty to f_t

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