Macroeconomic Dynamics, **21**, 2017, 918–946. Printed in the United States of America.
doi:10.1017/S1365100515000772

INTERNATIONAL SYNCHRONIZATION AND CHANGES IN LONG-TERM INFLATION UNCERTAINTY

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We investigate the international linkages of uncertainty associated with the long-term movements of inflation. In the first step, we establish that inflation uncertainty in the G7 is intertwined, and the degree of synchronization has increased during the recent two decades. We also document a rise in inflation uncertainty accompanying the global financial crisis. Based on a factor–structural vector autoregression, we provide evidence of a common international shock. We disclose that this shock is closely related to oil and commodity price uncertainty, and it explains large parts of the recent rise in inflation uncertainty. Moreover, increased synchronization can be explained by greater relative importance of this global shock. We also document that inflation uncertainty has become more stable, because domestic shocks translate less extensively into individual economies. This finding lends support to the "good policy" hypothesis.

Keywords: Inflation Uncertainty, Factor–Structural VAR, Stochastic Volatility

1. INTRODUCTION

It is well known that increased inflation uncertainty leads to economic costs. For instance, higher uncertainty about future inflation might trigger a disproportionate reallocation from nominal to real assets, and it makes nominal contracts involving wages and financial assets riskier [see, for instance, Fischer and Modigliani (1978); Bernanke and Mishkin (1997)].[1] Moreover, a strand of the literature stresses that higher inflation uncertainty is typically associated with higher inflation [Friedman (1977); Ball (1992); Cukierman and Meltzer (1986)]. 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]

We would like to thank Kai Carstensen, Helmut Herwartz, Matthias Hartmann, and participants at the 17th Spring Meeting of Young Economists, at the Annual Congress of the Verein für Socialpolitik 2012, at the CESifo Area Conference on Macro, Money and International Finance 2013, and at the Asian Meeting of the Econometric Society 2013 for very valuable comments and suggestions. Financial support from the German Research Foundation (Grant No. CA 833/2) is gratefully acknowledged. Address correspondence to: Steffen R. Henzel, Ifo Institute—Leibniz Institute for Economic Research at the University of Munich, Poschingerstrasse 5, 81679 Munich, Germany; e-mail: Henzel@ifo.de.

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This study sheds light on the international linkages of inflation uncertainty. We document the extent of comovement among the G7 and analyze the sources of international synchronization. To be able to distinguish between shocks common to all countries and spillovers of country-specific shocks, we estimate a factor-structural vector autoregression (FSVAR) model that allows decomposition into global shocks, spillovers stemming from other countries, and own shocks. Moreover, we split the sample and study the importance of different sources of fluctuations over time; i.e., we investigate whether the magnitude of shocks to inflation uncertainty has changed or whether the sensitivity toward these shocks has shifted.

In our study, the focus lies on inflation uncertainty associated with long-term movements of inflation. The reason is that most of the costs of inflation uncertainty, such as higher risk in long-term contracts, involve uncertainty over several years. In contrast, uncertainty about short-lived shifts in inflation, which are quickly offset in subsequent months, is likely of secondary importance for investment decisions or wage negotiations. Uncertainty about long-term inflation is thus the primary source of economic costs, which also makes long-term inflation uncertainty the major concern of policy makers [Ball and Cecchetti (1990); Cecchetti et al. (2007)].

A number of studies focus on common factors as a reason for business cycle synchronization [see, for instance, Stock and Watson (2005); Kose et al. (2008)]. Likewise, incomplete exchange rate adjustment and exposure to global shocks, such as 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. (2013) analyze international linkages of inflation between major industrialized countries. They provide evidence of increased comovement among the euro area countries as well as a rising correlation between the United States, Canada, and the euro area aggregate. We extend this literature by analyzing the degree and sources of synchronization of inflation uncertainty in the G7. We consider common shocks and spillover effects as possible explanations for synchronization and quantify the importance of each of these components.

Another strand of the literature documents a decline in the volatility of inflation in the United States since the mid-eighties [Stock and Watson (2007, 2010); Cogley et al. (2010); Canova and Ferroni (2012)]. Likewise, Cecchetti et al. (2007) demonstrate that the volatility of trend inflation in the G7 countries has decreased over time, which constitutes an "inflation stabilization" process. Moreover, Bataa et al. (2013, 2014) analyze the nature and timing of changes in international inflation uncertainty and document a decline in inflation uncertainty in the G7 in the mid-eighties.[3] To shed light on the sources of these changes, we quantify the role of the size of shocks governing long-term inflation uncertainty ("good or bad luck"), and we assess to what extent changes in the structure of the economy have altered the propagation of these shocks ("good or bad policy").

Our results can be summarized as follows. We provide evidence of synchronization among inflation uncertainty in the G7 economies. Moreover, inflation uncertainty tends to increase after the beginning of the 2000s, and the degree of synchronization among developed countries has increased over time. Controlling for spillover effects, we reveal a common shock that moves long-term inflation uncertainty in all G7 countries in the same direction. This global shock is closely related to oil and commodity price uncertainty. Moreover, it is capable of partly explaining the recent increase in long-term inflation uncertainty. Furthermore, we demonstrate that the relative importance of international shocks has increased over time, which provides an explanation for the higher degree of synchronization among the G7. Finally, we document that there has been a marked increase in the stability of inflation uncertainty. Our results suggest that "good policy" accounts for major parts of this decline, because domestic shocks translate less extensively into the individual economies.

The paper is organized as follows. We introduce our measure of long-term inflation uncertainty in Section 2. In Section 3, we examine the degree of synchronization among the G7 countries. The setup of the FSVAR model is explained in Section 4, whereas the empirical results of the FSVAR estimation are presented in Section 5. Section 6 concludes.

2. MEASURING INFLATION UNCERTAINTY

The measurement of unobserved uncertainty associated with the more persistent fluctuations of inflation poses a challenge. Ideally, a measure of long-term inflation uncertainty is derived from the subjective probability density forecasts for long-term inflation [see also Zarnowitz and Lambros (1987); Giordani and Söderlind (2003); Rich and Tracy (2010)]. However, consistent survey data for a longer time span including all G7 countries are not available at this present time. We resolve the problem using time-varying volatility, which we extract from the time-series dimension of data. This proceeding leads to an uncertainty measure that is consistently available for a long history and for all countries in our sample. Recent studies implement a model with stochastic volatility to measure time-varying uncertainty [see, for instance, Fernandez-Villaverde et al. (2011); Dovern et al. (2012)]. We follow this avenue because the stochastic volatility model—in contrast to a GARCH model—allows a separate innovation impinging on volatility [see, for instance, Fernandez-Villaverde and Rubio-Ramirez (2010)] . Furthermore, we need to separate uncertainty relating to the fluctuations of the more persistent component of inflation from uncertainty associated with transitory fluctuations. To achieve this decomposition, we employ the unobserved component stochastic volatility (UC-SV) model given by the following equations:

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

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

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

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

$$\left(\nu_{1,t} \quad \nu_{2,t}\right)' \sim N(0, \gamma I_2). \tag{5}$$

Specifically, the long-term component of inflation is given by $\bar{\pi}_t$, and it is governed by an idiosyncratic shock ϵ_t. The transitory component η_t captures short-lived fluctuations which are offset in the subsequent period. The innovations $\nu_{1,t}$ and $\nu_{2,t}$ inflate the volatility of the transitory and the long-term component, respectively.[4] An increase in the standard deviation $\sigma_{\epsilon,t}$ reflects that the permanent component of inflation is subject to larger changes, which translate into larger forecast errors about the long-term inflation outlook. This state-space model thus serves as a filtering device that enables us to extract a measure of long-term inflation uncertainty $\sigma_{\epsilon,t}$. The scalar parameter γ determines the smoothness of time-varying uncertainty.[5] The latter is obtained as a random walk to allow for permanent shifts.[6]

Using time-varying volatility as a proxy for unobserved uncertainty requires the assumption that the underlying model appropriately extracts the unforeseen shifts in trend inflation. Indeed, it has been shown that the UC-SV model delivers very accurate inflation forecasts [Stock and Watson (2007, 2010); Clark and Doh (2011)]. Although parsimonious, it enables the researcher to handle the nonlinearities involved in the inflation process; e.g., higher trend inflation is allowed to be accompanied by larger shocks. Also note that the measure is backward-looking, because it is based on in-sample information and delivers uncertainty prevailing at the time. Yet Grimme et al. (2014), among others, document considerable comovement between forward-looking survey-based measures of inflation uncertainty and stochastic volatility.

Our measurement approach complements the more structural treatment of uncertainty in, for instance, Bloom (2009) and Fernandez-Villaverde et al. (2011). The univariate UC-SV model measures uncertainty surrounding innovations in the permanent component of inflation. In a multivariate context, for instance, Cogley et al. (2010) measure stochastic volatility surrounding structural shocks involving restrictions from economic theory. Alternatively, Mumtaz and Zanetti (2013) propose a framework that enables them to measure time-varying structural shock volatility and the macroeconomic impact of changes in this measure, without making strong assumptions about the driving force behind such effects. Our measurement approach, however, reveals the patterns in the G7 countries without imposing further economic restrictions, and the empirical regularities we document provide the basis for a more structural treatment of global inflation uncertainty.

Uncertainty in the G7 (Canada, France, Germany, Italy, Japan, the United Kingdom, and the United States) is estimated by running the Gibbs sampler over the period 1960:M1–2012:M4.[7] We measure inflation as the annualized monthly percentage change in the Consumer Price Index (CPI) given by $1{,}200 \times \log(\mathrm{CPI}_t/\mathrm{CPI}_{t-1})$. Inflation series are retrieved from the OECD Main Economic Indicators (MEI) database ("CPI—All items") and are seasonally adjusted. The sources of the series are the respective national statistical agencies of countries. Finally, outliers in the data have been removed, most of which are attributable to announced changes in the value-added tax rate.[8] Figure 1 shows

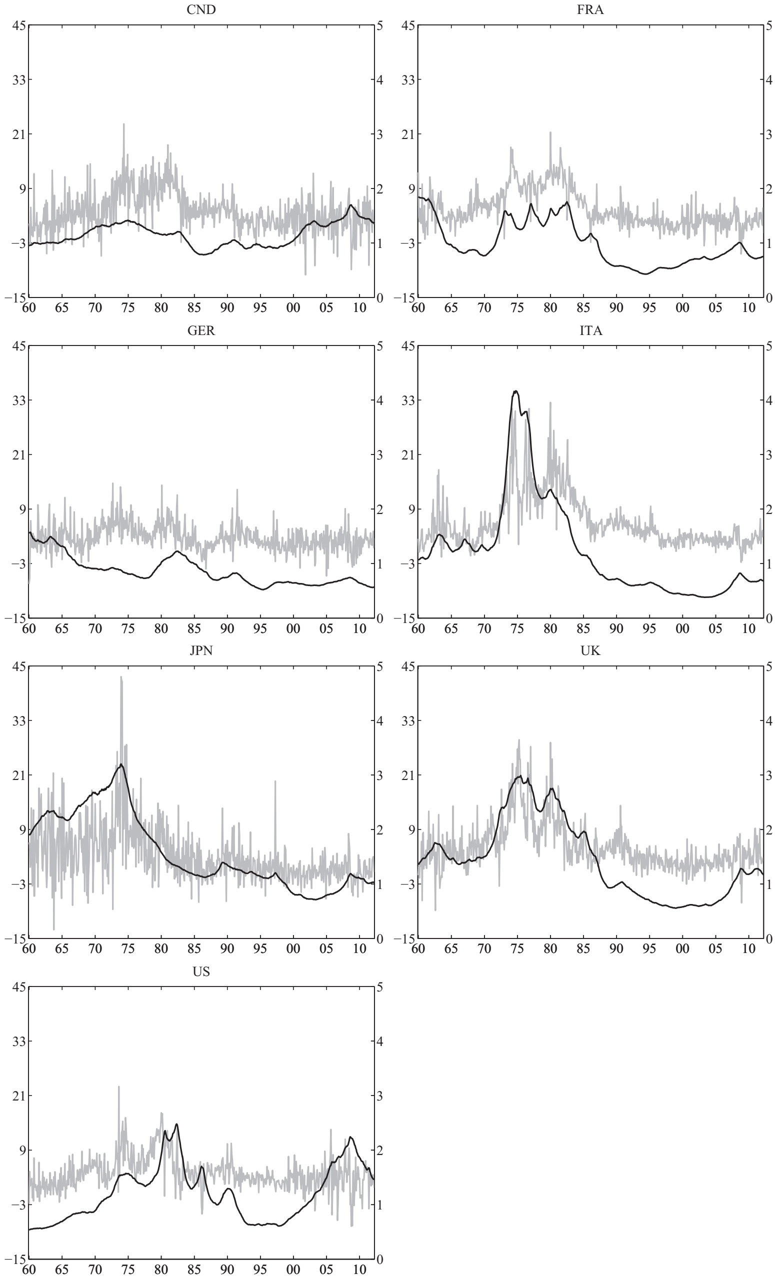

FIGURE 1. Inflation and long-term inflation uncertainty. The gray line represents actual inflation π_t (left-side axis); the dark line represents long-term inflation uncertainty measured by $\sigma_{\epsilon,t}$ as a percentage (right-side axis).

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TABLE 1. Pairwise correlations of inflation uncertainty

	CND	FRA	GER	ITA	JPN	UK
FRA	0.20***					
GER	0.14***	0.23***				
ITA	0.19***	0.18***	0.03			
JPN	0.21***	0.10**	−0.03	0.36***		
UK	0.15***	0.30***	0.23***	0.50***	0.20***	
US	0.33***	0.40***	0.32***	0.25***	0.09**	0.27***

Note: The entries indicate the correlations of country pairs. The uncertainty measures were differenced beforehand. Significance at the 1, 5, and 10% levels is indicated by ***, **, and *, respectively.

the measures of long-term inflation uncertainty along with the monthly inflation rates. A similar pattern emerges for all G7 economies. In light of the high inflation rates observed in the seventies, we measure a steady increase in uncertainty during this time. The upswing is followed by a marked reduction in uncertainty in the mid-eighties, which constitutes the process of 'inflation stabilization" [Cecchetti et al. (2007)]. However, during the last decade, uncertainty has risen 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) concerning this point].

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. Table 1 reveals that the pairwise correlations are positive and significant in the majority of cases, suggesting that inflation uncertainty in the G7 co-moves.

Synchronization among a group of countries may be assessed using cohesion, as proposed by Croux et al. (2001).[9] This measure is shown in the left part of Figure 2. We calculate cohesion for different frequencies on the interval $[0,\pi/4]$, that is, from long-term cycles with frequency zero up to the shortest cycle of 8 months. The shaded area indicates frequencies ranging from 1.5 to 8 years. In this range, cohesion is positive, suggesting that these lower frequencies contribute extensively to the co-movement of G7 countries. This result is also confirmed when we measure cohesion as a weighted average of G7 countries (dashed line in Figure 2).

Although 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. Following Bataa et al. (2014), it appears reasonable to split the sample at 1990, which is roughly in the middle of the sample. The right part of Figure 2 depicts cohesion calculated for the periods 1960–1989 and

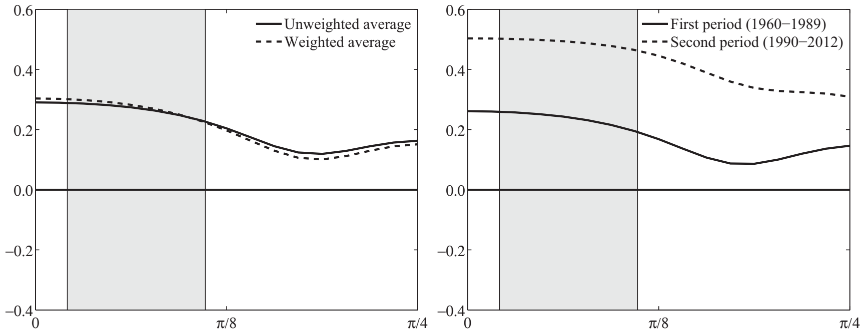

FIGURE 2. Cohesion of inflation uncertainty among the G7. The shaded area represents business cycle frequencies (8 to 1.5 years). The weighted average is calculated according to the country shares in aggregate GDP of the G7 (based on values in US$, constant prices and constant PPPs, OECD base year). The uncertainty measures were differenced beforehand. The spectra and the cospectrum involved are estimated using a Bartlett window with lag window size 12.

1990–2012, respectively. It becomes evident that cohesion increases considerably, i.e., inflation uncertainty co-moves more closely in the second subsample.[10]

4. THE FACTOR-STRUCTURAL VAR MODEL

The results presented in the previous section raise the question of why uncertainty is synchronized in the G7 economies. In general, there might be two reasons: common (global) shocks to inflation uncertainty and spillover effects from one country to another. To disentangle both channels, we rely on a factor-structural VAR (FSVAR) model of the following form:[11]

$$Y_t = A(L)Y_{t-1} + v_t, \tag{6}$$

$$v_t = \Lambda f_t + \zeta_t, \tag{7}$$

$$E(v_t v_t') = \Sigma_v, \tag{8}$$

$$E(f_t f_t') = \text{diag}\left(\sigma_{f_1}^2, ..., \sigma_{f_k}^2\right), \tag{9}$$

$$E(\zeta_t \zeta_t') = \text{diag}\left(\sigma_{\zeta_1}^2, ..., \sigma_{\zeta_7}^2\right). \tag{10}$$

Here, Y_t is a 7×1 vector stacking the demeaned uncertainty measures of the G7. The common shocks are captured by f_t, Λ is the $7 \times k$ matrix of factor loadings, and ζ_t denotes idiosyncratic shocks. By assumption, idiosyncratic shocks are uncorrelated with common shocks. We further restrict the lag polynomial $A(L)$ so that in each equation the lag length for the autoregressive part may differ from the number of lags of the other variables. To obtain a parsimonious specification, we choose the lag length by the BIC, which indicates four autoregressive lags and one lag of each of the remaining six variables in each equation. To ensure

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non-negative values of uncertainty, we take the log of $\sigma_{\epsilon,t}$. The FSVAR model is estimated using the maximum likelihood procedure following Stock and Watson (2005).[12]

According to (7), the error term of the FSVAR model is decomposed into country-specific idiosyncratic shocks and common shocks. The common shocks are identified by the assumption that they impact all countries immediately, whereas idiosyncratic shocks have an impact on other countries only via the autoregressive dynamics of the FSVAR model. We emphasize that in our case this assumption appears not overly restrictive. Because we use monthly data, spillovers may occur after one month. As a result, our model tends to attribute less explanatory power to common shocks than models based on quarterly data [see, for instance, Stock and Watson (2005); Carare and Mody (2010)]. If the number of common shocks is larger than one, these shocks need to be identified separately. A customary approach is to impose zero restrictions on the entries of Λ, the matrix of factor loadings [see, for instance, Stock and Watson (2005); Gorodnichenko (2006)]. 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 (United States), and the third factor has zero impact on both last-ordered countries (United Kingdom and United States).

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 seven idiosyncratic shocks, whereas the alternative states 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. The test supports one common factor, as the null of $k = 1$ cannot be rejected at the 5% level.[13]

Table 2 also contains the LR test for two subsamples. In the second subsample (1990–2012), the way toward EMU might have affected inflation dynamics in three of the G7 countries. To analyze whether there is an additional EMU-specific factor, we conduct the LR test for the second subsample. The result is shown in the right panel of Table 2. Because we cannot reject the restrictions involved by one common shock for the more recent period, it appears that there is no EMU factor in addition to the common factor. Apparently, we cannot reject one common factor for the first period (1960–1989) either.

5. EMPIRICAL RESULTS

This section presents results of the FSVAR estimation. We first investigate the response of individual countries to the common shock and relate it to other global uncertainty measures. Second, we analyze the importance of the different types of shocks and estimate how international components contribute to the recent increase in inflation uncertainty. Third, we assess whether changes in the shock size or in the propagation of shocks may account for increased synchronization

TABLE 2. Testing for the number of common factors in the FSVAR model

		1960–2012			1960–1989			1990–2012		
k	d.f.	log L (10^4)	LR stat.	p-value	log L (10^4)	LR stat.	p-value	log L (10^4)	LR stat.	p-value
0	–	20.4455	–	–	11.6409	–	–	8.8354	–	–
1	14	20.4360	19.03	0.16	11.6306	20.58	0.11	8.8269	16.83	0.27
2	8	20.4414	8.28	0.41	11.6373	7.15	0.52	8.8335	3.65	0.89
3	3	20.4450	1.14	0.77	11.6406	0.59	0.90	8.8344	2.01	0.57

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.

and whether and how the dynamics of long-term inflation uncertainty has changed during the process of inflation stabilization.

5.1. The Common Shock

To see whether the common shock f_t qualifies for a global driver of inflation uncertainty, we calculate the response to a surprise increase in f_t of one standard deviation. Figure 3 displays the impulse response functions of the individual countries. A surprise innovation in the international factor significantly shifts inflation uncertainty upward in all countries.[14] The impulse response follows a hump-shaped pattern, with a strong reaction in France, Italy, the United Kingdom, and the United States and a less pronounced increase in Canada, Germany, and Japan. Because the common shock uniformly drives uncertainty in the G7 economies in the same direction, it qualifies as a global driver of inflation uncertainty and thus provides an explanation for the synchronization among the G7.[15]

Figure 3 also reveals that the response to the common shock is positive in both subsamples. The interpretation of the common shock as a global driver of inflation uncertainty thus remains valid when subsamples are considered. When compared with the first subsample, the European countries appear to experience a dampened response in the second subsample. The North American countries seem to display a more pronounced increase, which appears reasonable because the global financial crisis affected those countries more heavily. Yet it should be noted that the observed differences are not statistically significant. For Japan, there is virtually no change when subsamples are considered.

5.2. Relation of Global Shock to Global Uncertainty Measures

In the following we aim to provide an economic interpretation of the global shock to long-term inflation uncertainty. Although the FSVAR method delivers shocks that are orthogonal, we have to rely on indirect evidence for an economic interpretation of these shocks. A possible driver of international inflation uncertainty is the uncertainty associated with prices of goods that are traded all over the world at a common price, and that 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 global shocks. We can infer whether the global shock f_t and any other measure of uncertainty are related by estimating the following regression:

$$\mathrm{unc}_t^i = c_0 + \sum_{j=0}^{J} \phi_j^i f_{t-j} + \varepsilon_t^i. \tag{11}$$

Here, unc_t^i represents a measure of uncertainty, and ε_t^i is the respective regression residual. Because f_t is an orthogonal white noise process, it is exogenous in such a

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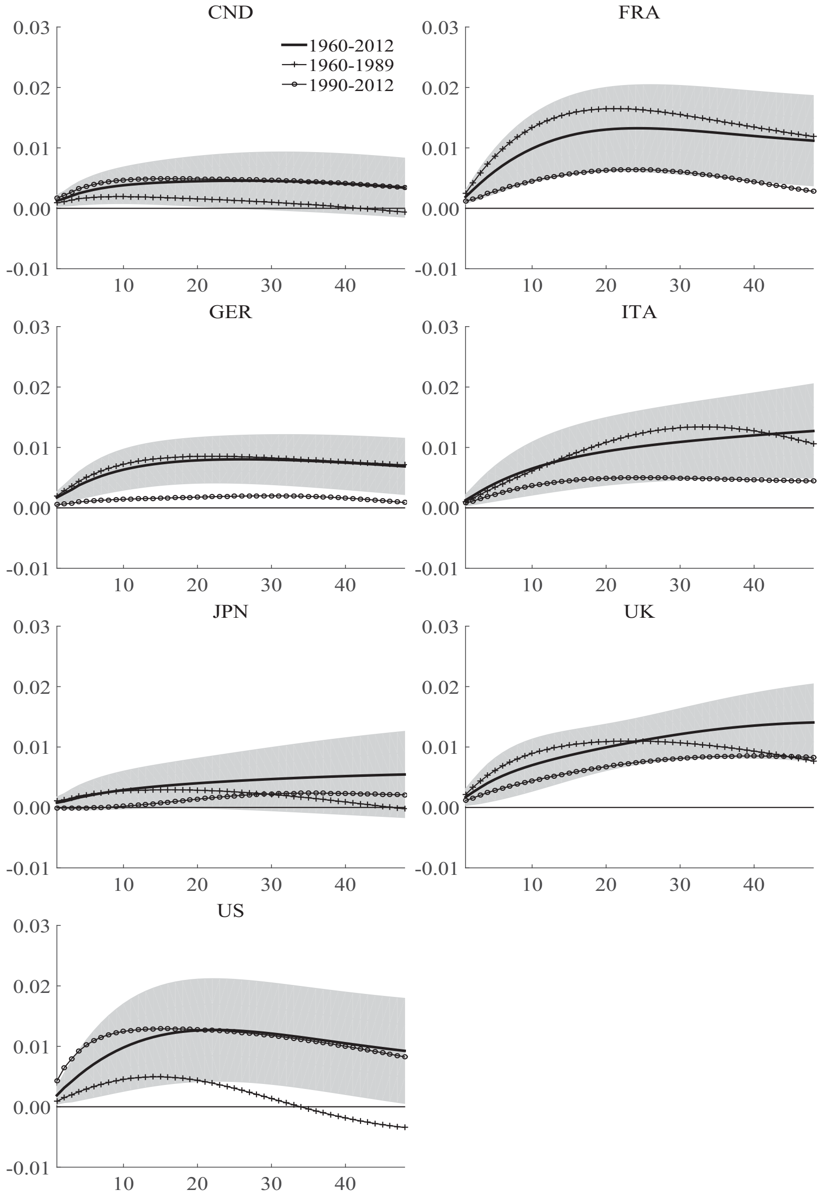

FIGURE 3. Response of inflation uncertainty to the common shock. Black lines show the response of inflation uncertainty to a one-standard-deviation increase in the common shock for the whole sample (1960–2012). Shaded areas are 95% bootstrap confidence bands. Crossed lines show the response of inflation uncertainty for the sample 1960–1989; bullet markers indicate the sample 1990–2012.

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regression. The estimated coefficients ϕ_j^i provide the moving average representation of the dynamic relation between unc_t^i and the global inflation uncertainty; i.e., the cumulative coefficients are the impulse response function of unc_t^i following a 1% increase in the exogenous shock f_t [see, for instance, Kilian (2009); Romer and Romer (2010)]. We consider $J = 24$ 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 because 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 because 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 the upper panel of Figure 4 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 global shock to inflation uncertainty. Notably, a positive global shock foreshadows increases in oil and commodity price uncertainty.[16]

Alternatively, global uncertainty about fluctuations in the exchange rate might be closely related to the common shock driving long-term inflation uncertainty.[17] In the lower panel of Figure 4 there appears to be no significant relation between the common shock and exchange rate uncertainty. Likewise, uncertainty perceived on financial markets might be associated with the global shock. We also run regression (11) for financial market uncertainty.[18] The global shock and financial market uncertainty are unrelated. In sum, our results provide evidence that f_t may be interpreted as a shock to commodity price uncertainty, which shows up as a global shock to inflation uncertainty in the G7.[19]

5.3. Importance of Global Shock

In the following, we assess the importance of the respective shocks driving long-term inflation uncertainty with a focus on the international component (global shock and spillovers). Because ζ_t and f_t in the FSVAR model are uncorrelated by assumption, total forecast error variance for each country can be decomposed into the global shock, the own shock and spillovers received from the remaining six countries. Based on the FSVAR estimation, Table 3 displays the forecast error standard deviation and the contribution of the different types of shocks to the forecast error variance of inflation uncertainty at forecast horizons up to 48 months.

Overall, the results of the variance decomposition suggest that there are non-negligible international linkages. For all countries, the proportion of the domestic

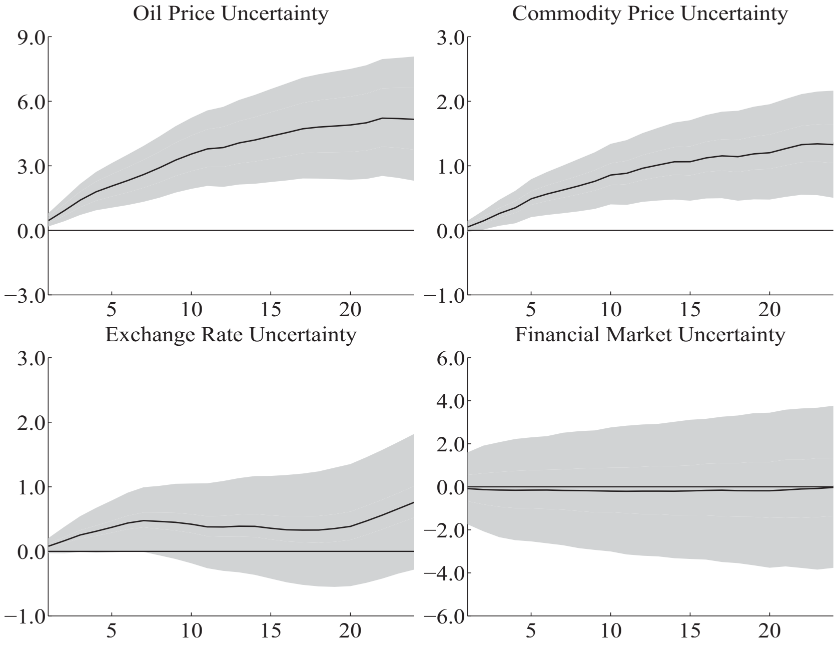

FIGURE 4. Uncertainty measures and the global shock f_t. The solid line represents the response of unc_t^i to a 1% increase in f_t. Shaded areas represent the 95% error bands, which are obtained by a block bootstrap using a block size of 12 and 20,000 replications. The different uncertainty measures unc_t^i are WTI oil price uncertainty (upper left), CRB commodity price uncertainty (upper right), exchange rate uncertainty (lower left), and financial market uncertainty (lower right). Because of data availability, the estimation sample starts in 1979M6 for oil price uncertainty, in 1962M8 for financial market uncertainty, and in 1960M1 for the remaining series.

component declines with the forecast horizon in favor of the international component, i.e., in favor of the global shock or spillovers. Indeed, the global shock has a noticeable impact on the fluctuations of long-term inflation uncertainty. For the euro area countries and the United Kingdom it captures between 12% and 19% of the variance, and for the United States and Canada it contributes 7% and 10%, respectively, whereas the contribution of the global shock is the smallest in Japan of all countries. The proportion of spillover increases with the forecast horizon. This share amounts to 30% in Italy. The United States receives little spillover from abroad. Likewise, the share of spillover in Germany is also comparatively low. These findings are in line with the fact that the Federal Reserve Bank and the German Bundesbank have conducted monetary policy largely independently during most of the sample. Domestic shocks are particularly important in Japan, which probably reflects the turbulent situation of the Japanese economy during the nineties and early 2000s.

Table 3. Variance decomposition into global shock, spillover, and own shocks

			Fraction of FEV due to						Fraction of FEV due to		
	Hor.	FE std.	Global	Spillover	Own		Hor.	FE std.	Global	Spillover	Own
CND	1	0.00	6.36	0.00	93.64	FRA	1	0.01	9.03	0.00	90.97
	12	0.27	6.58	0.69	92.73		12	0.50	9.60	0.19	90.22
	24	0.90	6.95	5.22	87.83		24	1.88	10.54	1.40	88.06
	48	2.86	6.93	20.54	72.52		48	5.83	12.57	6.62	80.81
GER	1	0.01	10.66	0.00	89.34	ITA	1	0.01	4.64	0.00	95.36
	12	0.32	10.99	0.17	88.84		12	0.44	5.31	0.37	94.32
	24	1.14	11.72	1.44	86.83		24	1.57	6.89	3.69	89.41
	48	3.55	13.30	7.94	78.76		48	4.67	11.60	30.22	58.17
JPN	1	0.01	2.37	0.00	97.63	UK	1	0.01	8.54	0.00	91.46
	12	0.31	2.37	0.19	97.44		12	0.37	9.99	0.37	89.64
	24	1.19	2.39	1.54	96.07		24	1.24	13.10	3.65	83.25
	48	4.43	2.44	7.77	89.79		48	4.01	18.98	22.81	58.21
US	1	0.01	8.35	0.00	91.65						
	12	0.53	8.66	0.17	91.17						
	24	2.00	9.10	1.34	89.56						
	48	6.32	9.79	7.37	82.84						

Note: This table displays the standard deviation and variance decomposition of inflation uncertainty forecast errors at the 1-, 12-, 24-, and 48-month horizons. The F.E. standard deviation is reported in percentage points. Estimation based on an FSVAR model with one common factor, four autoregressive lags, and one lag for the remaining six countries in each equation.

Notably, these results confirm our identification strategy. Given the predominant role of the United States, markets around the globe may react quickly, i.e., within one month, to changes in U.S. monetary policy, for instance. If this were the case, a pure U.S. shock might wrongly be identified as a global shock. We show, however, in Table 3 that the common shock is not predominantly associated with the United States, because the global shock does not explain more of the U.S. forecast error variance than that of most other countries.

As discussed in Section 4, the way toward EMU might have a bearing on the results in the second half of the sample. Although Table 2 reveals that there is only one common (i.e., global) shock in the second subsample, as a further robustness check, we may yet impose an additional EMU shock. We thus include a second shock that impacts only France, Germany, and Italy and estimate the FSVAR model for the years 1999–2012, i.e., the EMU period. However, the second shock has explanatory power almost exclusively for Germany. Moreover, this shock affects neither the interpretation nor the importance of the global factor.[20]

We can also assess the importance of international shocks using counterfactual simulations. In the following, we focus on the recent increase in inflation uncertainty documented in Section 2, which suggests that the period of inflation stabilization has come to an end. From a monetary policy perspective it is important to know whether this increase in inflation uncertainty has its origins in international or domestic shocks, and the analysis enables us to reveal which shock contributes to the recent increase. Figure 5 shows the measure of long-term inflation uncertainty along with the time path that would have been observed if only the global shock operated on long-term inflation uncertainty, i.e., if we shut down the remaining seven shocks. It appears that the dynamics of inflation uncertainty are to a considerable extent governed by the global shock. Particularly, the recent hike that accompanied the global financial crisis starting in 2007 is traceable to this shock. Given that the global component is closely related to oil and commodity price uncertainty, the analysis suggests that major parts of the recent increase come from the oil and commodity markets. Figure 5 also depicts the dynamics of inflation uncertainty when only spillovers from abroad are present. In Italy and in the United Kingdom, spillovers from other countries are capable of partly explaining the increase in long-term inflation uncertainty during the global financial crisis, whereas in the remaining countries spillovers have no considerable impact in this period.

5.4. Changes in the Dynamics of International Inflation Uncertainty

Results in Section 3 point to a greater degree of synchronization in the second subsample (1990–2012). Moreover, our sample comprises the inflation stabilization period, which might be accompanied by changes in the dynamics of inflation uncertainty. In the following, we trace the importance of different sources of fluctuations over time. We split the sample roughly in half and consider the more turbulent years 1960–1989 and the period of low and stable inflation 1990–2012.

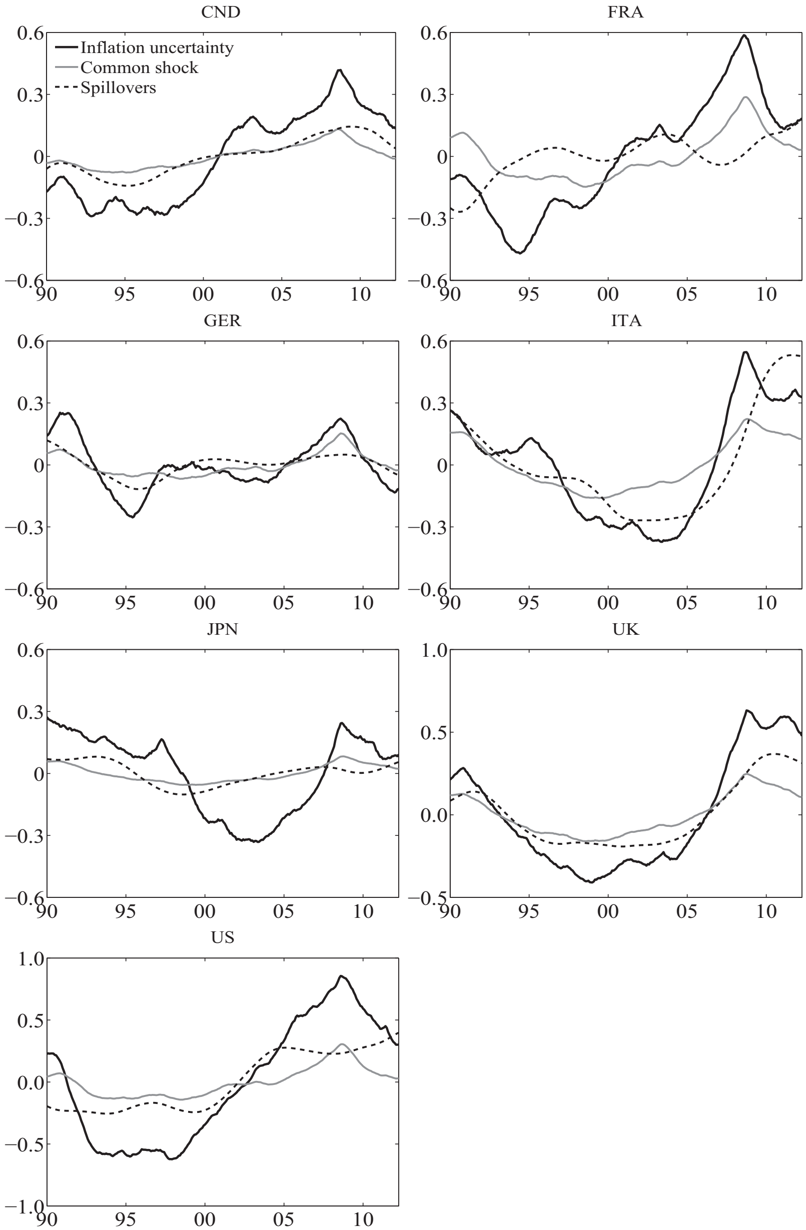

FIGURE 5. Counterfactual dynamics of long-term inflation uncertainty. The black line shows long-term inflation uncertainty, the gray line represents the counterfactual time path when only the common shock is active, and the dashed line depicts the dynamics resulting from spillovers. All series are demeaned.

https://doi.org/10.1017/S1365100515000772 Published online by Cambridge University Press

In general, changes in the variability of long-term inflation uncertainty may be the result of either a change in the size of shocks ("good or bad luck") or structural changes in the economy ("good or bad policy"), and we decompose the change in total forecast error variance accordingly.

Let V_p denote the variance of the forecast error, where $p = 1, 2$ refers to the first and second subsamples, respectively. For notational simplicity, we suppress the dependence on the forecast horizon and the country. Because the FSVAR model incorporates eight sources of variation (one international shock and seven 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 subsample 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: $V_{p,j} = a_{pj}\sigma_{pj}^2$, where a_{pj} is given by the cumulated squared impulse responses to a standardized (unit) shock j. σ_{pj}^2 denotes the variance of shock j in subsample p. For each shock j, the change in the contribution to the total variance can be expressed as

$$V_{2j} - V_{1j} = \left(\frac{a_{1j} + a_{2j}}{2}\right)\left(\sigma_{2j}^2 - \sigma_{1j}^2\right) + \left(\frac{\sigma_{1j}^2 + \sigma_{2j}^2}{2}\right)(a_{2j} - a_{1j}). \qquad \textbf{(12)}$$

The first term on the right-hand side in (12) refers to the contribution from the change in the shock size, whereas the second term refers to the contribution from the change in the impulse response function. Note that idiosyncratic shocks originating abroad are summarized as "spillover." Table 4 shows total forecast error variance for the first part of the sample (1960–1989) in column (1) and for the second part (1990–2012) in column (2). As indicated by column (3), the variance has significantly decreased in the majority of countries (FRA, GER, ITA, UK, US); i.e., the process of long-term inflation uncertainty has become more stable and its predictability has increased. In the remaining countries (CND, JPN), changes are not significant.

Columns (4) to (11) of Table 4 report the decomposition of changes in the 12-months-ahead forecast error variance of inflation uncertainty into the distinct sources of variation. Column (4) reveals that the magnitude of the common shock has increased over time, whereas domestic shocks reported in column (6) contribute negatively in almost all countries. Column (5) shows that foreign shocks (spillovers) contribute negatively in all countries. This decline is, however, small compared with the increase in the common shock. In sum, international shocks have gained importance relative to domestic shocks. Column (8) shows that the common shock has a larger influence on Canada and the United States, whereas the propagation of those shocks is dampened in the remaining countries. Contributions of changes in the propagation of shocks from abroad are rather heterogeneous across countries [see column (9)]. Column (10) indicates that domestic shocks are transmitted less extensively into all countries but Canada. In sum, domestic shocks lose importance relative to the global shock, whereas changes in spillovers are of

TABLE 4. Decomposition of change in forecast error variance

	Total variances			Contribution of change in shock size				Contribution of change in impulse responses			
	(1) 1960–1989	(2) 1990–2012	(3) Change	(4) Common	(5) Spillover	(6) Domestic	(7) Total	(8) Common	(9) Spillover	(10) Domestic	(11) Total
CND	4.328***	5.837***	1.508	0.127	−0.029	0.743	0.841	0.465	0.062	0.140	0.667
	(0.693)	(1.107)	(1.297)	(0.225)	(0.040)	(0.806)	(0.856)	(0.825)	(0.141)	(1.037)	(1.380)
FRA	30.748***	6.359***	−24.389***	0.877	−0.038	−9.660***	−8.820***	−4.904	−0.164	−10.501***	−15.569***
	(4.583)	(1.307)	(4.760)	(1.084)	(0.080)	(2.983)	(3.052)	(3.415)	(0.323)	(3.172)	(4.440)
GER	11.642***	4.098***	−7.544***	0.281	−0.022	−1.713	−1.454	−1.741	0.296	−4.645***	−6.090***
	(2.053)	(0.691)	(2.174)	(0.411)	(0.049)	(1.366)	(1.412)	(1.699)	(0.209)	(1.676)	(2.201)
ITA	21.181***	10.713***	−10.467***	0.187	−0.016	−6.564***	−6.393***	−0.662	−0.099	−3.312	−4.074
	(3.445)	(2.130)	(4.011)	(0.409)	(0.057)	(2.272)	(2.346)	(1.618)	(0.215)	(3.241)	(3.742)
JPN	5.121***	6.938***	1.817	0.046	−0.046	1.872**	1.873**	−0.303	0.307	−0.060	−0.056
	(0.755)	(1.354)	(1.543)	(0.128)	(0.055)	(0.789)	(0.843)	(0.533)	(0.218)	(1.297)	(1.382)
UK	16.866***	7.007***	−9.858***	0.470	−0.029	−1.497	−1.057	−2.210	0.059	−6.650***	−8.802**
	(2.751)	(1.270)	(3.041)	(0.879)	(0.042)	(2.784)	(2.600)	(3.213)	(0.171)	(2.386)	(3.624)
US	27.110***	12.040***	−15.071***	0.853	−0.049	−10.893**	−10.089**	4.389	−0.240	−9.130**	−4.981
	(4.887)	(2.090)	(5.311)	(1.149)	(0.080)	(5.360)	(4.764)	(3.163)	(0.302)	(3.861)	(5.781)

Note: Columns (1) and (2) show the 12-months-ahead forecast error variance decomposition of inflation uncertainty for two subsamples and the difference between the two subsamples based on an FSVAR estimation with one common factor, four autoregressive lags, and one lag of the remaining six countries in each equation. Changes are reported in column (3). Columns (4) to (6) report the change in the magnitude of the common shock, country-specific shocks from abroad (spillover), and domestic shocks. Column (7) shows the total contribution of the change in shock size. Columns (8) to (10) report contributions of change in propagation of common shocks, country-specific shocks from abroad (spillover), and domestic shocks into the domestic economies. The sum of these contributions is reported in column (11). The values are multiplied by 100, and bootstrap standard errors are reported in parentheses.

secondary importance. Increased synchronization is thus the result of domestic shocks losing importance relative to the common shock.

In sum, there has been a reduction in the magnitude of shocks, which is statistically significant for France, Italy, and the United States [see column (7)]. Such a result suggests that "good luck" has contributed to the decline of the volatility of inflation uncertainty in these countries. As reported in column (11), changes in the impulse responses significantly contributed to the overall decline in variance as well. The reason is that the sensitivity toward domestic shocks has decreased in all countries except Canada. It appears that "good policy" is responsible for large parts of the decline in the volatility of inflation uncertainty. Changes in the propagation mechanism of shocks to inflation uncertainty have apparently contributed to a moderation in long-term inflation uncertainty which parallels the process of inflation stabilization.[21]

Given the relatively high importance of "good policy," the question arises of which policy led to a stabilization of long-term inflation uncertainty. One policy area that underwent major changes in the last two decades is the field of monetary policy. Nowadays, there seems to be a better understanding of how to implement monetary policy, with central banks being more responsive to inflationary shocks [see, for instance, Clarida et al. (2000); Summers (2005); Cecchetti et al. (2006)]. Notably, Cecchetti et al. (2007) document that the observed process of inflation stabilization in the G7 has come along with a common shift in central banks' behavior, from an accommodative to a more responsive stance. This shift in the conduct of monetary policy has been accompanied by a number of institutional changes concerning monetary authorities in the G7. During the 1990s, France, Italy, Japan, and the United Kingdom implemented major legislative reforms, which enhanced central bank independence. Likewise, indices of political and economic autonomy have generally risen in the G7 countries from the first half of our sample to the second half [see Acemoglu et al. (2008); Arnone et al. (2009)]. Moreover, during the last two decades, all countries in our sample have introduced inflation-targeting strategies, providing a strong nominal anchor for inflation expectations [Mishkin and Schmidt-Hebbel (2007)].[22] Overall, our results suggest that these changes in the field of monetary policy not only reduced inflation uncertainty in the last two decades but also contributed to a stabilization of inflation uncertainty.

6. CONCLUDING REMARKS

Our study provides insight into the international linkages of inflation uncertainty. The results can be summarized as follows. First, we find evidence of synchronization among uncertainty surrounding the long-term movements of inflation in the G7. We show that the degree of synchronization has increased during the most recent two decades. Second, in a FSVAR framework, we reveal one common shock that moves national long-term inflation uncertainty in all countries in the same direction. We find that this global shock is closely related to oil and commodity price uncertainty. Third, our results provide support for the claim that international factors are economically important for domestic inflation uncertainty. Fourth, we consider two subsamples and reveal that higher connectedness of inflation uncertainty among the G7 is traceable to an increase in the relative importance of the global shock. Whereas domestic sources lose importance, worldwide uncertainty about oil and commodity prices becomes more relevant. Finally, we document that changes in the propagation mechanism of shocks to inflation uncertainty in the G7 increased the stability of inflation uncertainty. The main channel are domestic shocks that translate less extensively into the individual economies. This finding supports the hypothesis of "good policy."

As stressed by Cecchetti et al. (2007), the main candidate for inflation stabilization in the G7 is changes in central banking practices across the G7. This also provides a candidate explanation for the global "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 that credibly fight inflation. Accepting higher inflation—as recently called for to deal with the problem of excessive debt—may bring about the additional cost of higher worldwide inflation uncertainty via international linkages. Moreover, as observed during the recent global financial crisis, international oil and commodity price movements tend to counteract the inflation stabilization process.

NOTES

1. A growing amount of literature documents the potential effects of uncertainty on the real economy. See Bloom (2009), Bachmann et al. (2013), Baker et al. (2013), and Henzel and Rengel (2014), among others.

2. Consequently, a large number of empirical studies analyze 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), Caporale et al. (2012), and Hartmann and Roestel (2013).

3. Bataa et al. (2014) document that in Canada, in the United States, and (to a lesser extent) in the euro area, the decline is only temporary and the volatility of inflation shocks began to rise again at the beginning of the 2000s. For the euro area see also Hartmann and Herwartz (2014).

4. Our estimates reveal that the volatility of short-term movements is indeed fluctuating over time. It is thus advisable to allow for stochastic volatility of both long-term and short-term movements.

5. Stock and Watson (2007) calibrate this parameter to $\gamma = 0.20$ for quarterly inflation rates. Because we have monthly data, which usually carry more noise, we use $\gamma = 0.2/3 = 0.07$.

6. Alternatively, for instance, Grassi and Proietti (2010) estimate a first-order autoregressive process for volatility. However, they reach the conclusion that the autoregressive coefficient is very close to one.

7. Calculations are based on the replication files of Stock and Watson (2007), which are available from Mark W. Watson's website: www.princeton.edu/ mwatson/publi.html.

8. See Appendix A for a detailed description of outlier adjustment.

9. See Appendix B for a detailed description of how cohesion is calculated.

10. We test whether the difference between the two subsamples is statistically significant. To be compatible with the cohesion measure, we calculate changes in the bivariate correlations for the bandpass-filtered version of inflation uncertainty. Appendix C shows the difference in pairwise

correlations between the subsamples 1960–1989 and 1990–2012. Evidently, the majority of pairwise correlations have increased significantly.

11. The FSVAR model was originally used by Altonji and Ham (1990), Norrbin and Schlagenhauf (1996), and Clark and Shin (2000) to model regional spillovers. For an application to international linkages, see Stock and Watson (2005), Lahiri and Isiklar (2009), and Carare and Mody (2010), amongst others.

12. During estimation, we check whether the VAR in Equation (6) is stable, i.e., whether the maximum eigenvalue is smaller than 1, and find no violation of the stability condition.

13. We also performed the test estimating an unrestricted FSVAR model with 12 lags in each equation in Appendix D. This test also indicates one common shock.

14. In Appendix D we present results for a FSVAR model with 12 lags. The choice of this greater lag length does not affect the results.

15. In Appendix E, we analyze how the global shock differs from a shock originating in the United States. The U.S. shock is not capable of explaining the synchronization among the G7.

16. To analyze whether the result is driven by a few extreme observations in the WTI oil price, we replace observations that deviate more than six times the interquartile range from the local mean (Feb/Mar/Aug 1986, Aug/Sep 1990, Oct/Nov/Dec 2008) with the means from the six neighboring observations and reestimate equation (11). The relationship between the common shock to inflation uncertainty and oil and commodity price uncertainty remains significant.

17. We measure global uncertainty in FX markets based on bilateral monthly exchange rates of the G7 countries. Specifically, we calculate individual exchange rate uncertainty measures for each country using the UC-SV model introduced in Section 2. The first principal component across individual uncertainty measures in the G7 serves as a proxy for global exchange rate uncertainty. Our results are, however, robust to using only uncertainty associated with the US$–EUR exchange rate.

18. Financial market uncertainty is measured by 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 time series of global financial market uncertainty beginning in 1962M8. For the estimates in the present paper, we have updated Bloom's series to 2012M4. The VIX/VXO has also been used as a proxy for global financial market uncertainty by Milesi-Ferretti and Tille (2011), Forbes and Warnock (2012), and Carrière-Swallow and Céspedes (2013), among others.

19. This interpretation of the common shock is also supported by the observation that countries with a larger share of fuel in the CPI basket tend to deliver a larger response to the common shock in Figure 3.

20. This result squares well with previous studies on global inflation [see, for instance, Ciccarelli and Mojon (2010); Mumtaz and Surico (2012)]. Detailed results are available upon request.

21. The second subsample (1990–2012) also comprises the global financial crisis, which led to increased volatility in important macroeconomic aggregates, suggesting that the Great Moderation has come to an end. Likewise, it might be the case that there has been a break in the process of long-term inflation uncertainty. We analyze this issue in Appendix F. There are, however, no signs of a significant increase in forecast error variance of long-term inflation uncertainty during the crisis period (2007–2012).

22. Among the G7, Canada and the United Kingdom 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." Since the beginning of 2012, the United States and Japan have also communicated inflation targets.

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APPENDIX A: OUTLIER ADJUSTMENT

Table A.1 summarizes the adjustment of outliers. First, we identify outliers that are traceable to policy changes; in most cases, we identify an increase in the value-added tax rate. Second, a number of extreme observations are removed that are associated with outstanding events. The outliers in France in 1965M6 and 1965M7 are due to a shift in the level series of the CPI. Finally, we follow Stock and Watson (2003) and refer to an outlier if an observation deviates from the local mean by more than six times the interquartile range. These outliers are marked with an asterisk in Table A.1. All outliers are replaced with the means of the six adjacent observations.

TABLE A.1. Adjustment of inflation outliers

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%				

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APPENDIX B: CALCULATION OF COHESION

In the bivariate case, dynamic correlation between two variables x and y is defined as

$$\rho_{xy}(\lambda) = \frac{C_{xy}(\lambda)}{\sqrt{S_x(\lambda)S_x(\lambda)}}, \tag{B.1}$$

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 [see Croux et al. (2001)]. The spectra and the cospectrum involved are estimated using a Bartlett window with lag window size 12. The frequency λ is inversely related to the number of periods per cycle, $p = \frac{2\pi}{\lambda}$. Given monthly data, a frequency of $\frac{\pi}{4}$, for example, corresponds to a cycle of 8 months. 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:

$$\text{coh}_X(\lambda) = \frac{\sum_{i \neq j} w_i w_j \rho_{x_i x_j}(\lambda)}{\sum_{i \neq j} w_i w_j}, \tag{B.2}$$

where X denotes a vector of variables with entries x_i, and w_i denotes the respective weight of country i. We consider equal weights ($w_i = 1$) and weights according to the country's share in the aggregate GDP of the G7 economies.

APPENDIX C: TESTING FOR CHANGES IN CORRELATIONS AMONG COUNTRY PAIRS: DIFFERENCES IN PAIRWISE CORRELATIONS OF INFLATION UNCERTAINTY

	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 differences in correlation between the two subsamples. 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 that extracts business cycle frequencies (1.5 to 8 years).

APPENDIX D: ROBUSTNESS TESTING FOR THE NUMBER OF COMMON FACTORS IN THE FSVAR MODEL WITH 12 LAGS

k	$\log L$ (10^4)	d.f.	LR Stat.	p-value
0	20.5003	–	–	–
1	20.4897	14	21.30	0.094
2	20.4966	8	7.38	0.496
3	20.4991	3	2.48	0.479

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.

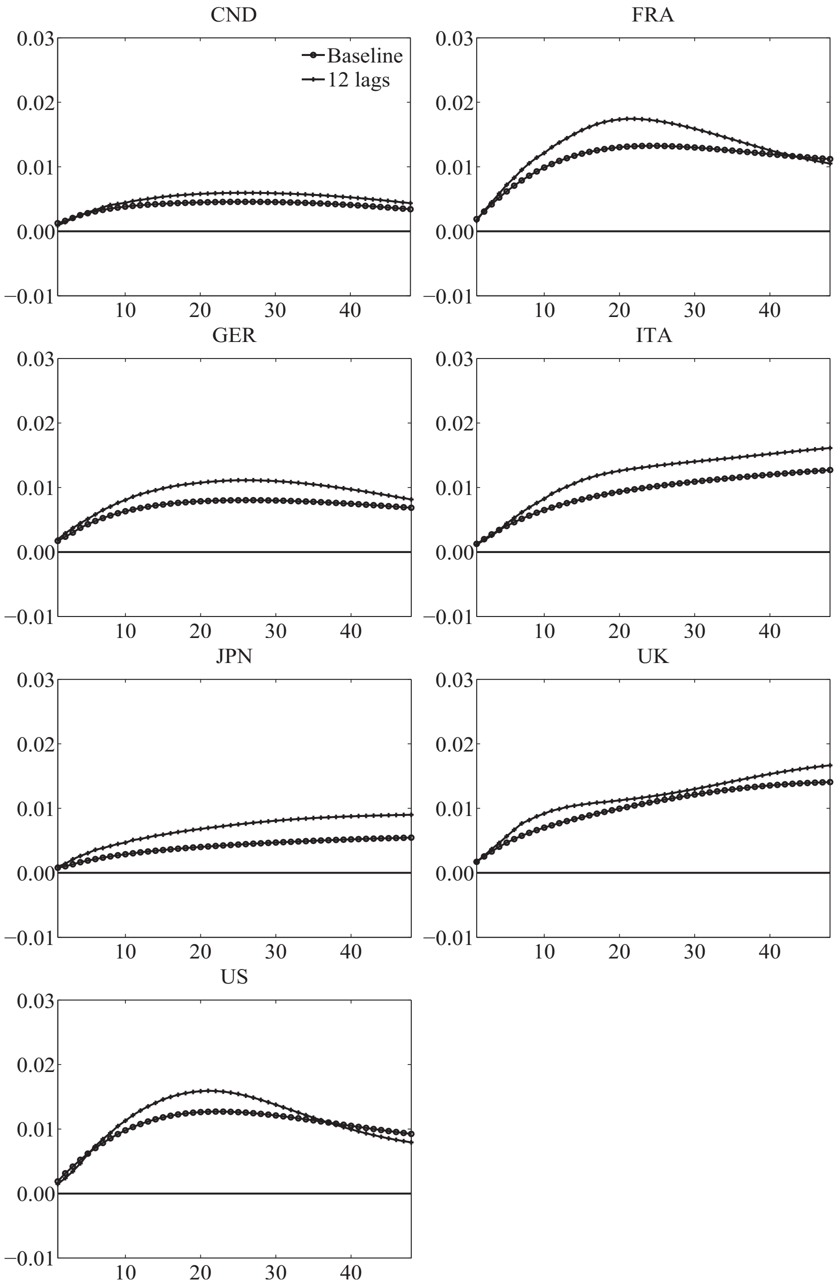

FIGURE D.1. Response of inflation uncertainty to the common shock in the FSVAR model with 12 lags. Bullet markers represent the response of inflation uncertainty in the baseline FSVAR model, and the crossed lines show the response of inflation uncertainty in an FSVAR model with 12 lags to a one-standard-deviation increase in the common shock.

https://doi.org/10.1017/S1365100515000772 Published online by Cambridge University Press

APPENDIX E: A COUNTRY-SPECIFIC SHOCK IN THE UNITED STATES

Also of interest is whether the global shock may be distinguished from a shock originating in the United States, and whether shifts in the United States are capable of explaining international synchronization. The impulse responses to a surprise innovation in U.S. inflation uncertainty are shown in Figure E.1. In contrast to the global shock, the response to a U.S. shock is insignificant in all countries except for Canada and the United Kingdom. In sum, there are marked differences between a country-specific shock originating in the United States and a global shock. Most notably, the U.S. shock is not capable of explaining the synchronization among the G7.

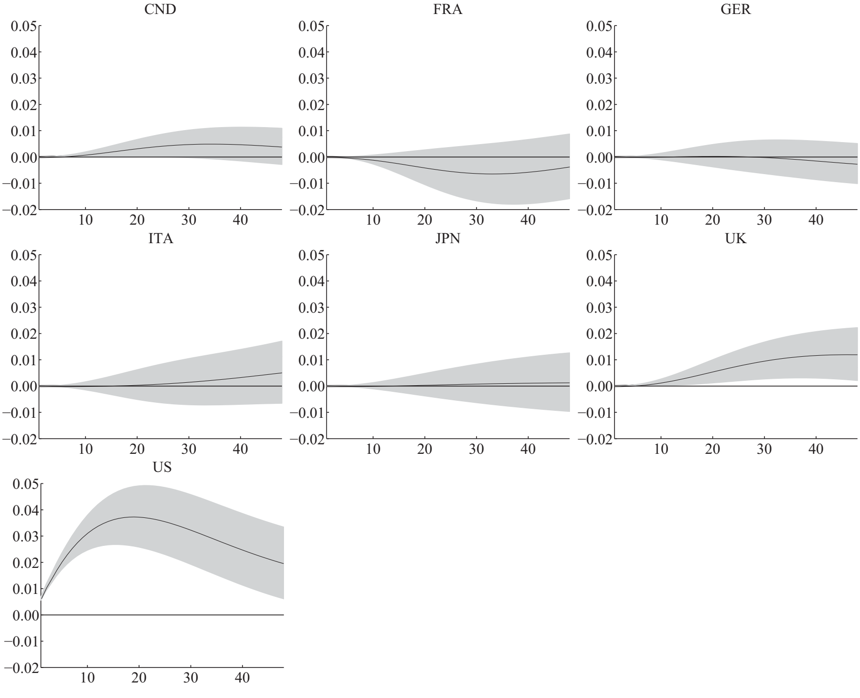

FIGURE E.1. Response of inflation uncertainty to a U.S. shock. The bold black line represents the response of inflation uncertainty in the respective country to a one-standard-deviation shock to inflation uncertainty originating in the United States. The shaded area represents the 95% bootstrap confidence band.

https://doi.org/10.1017/S1365100515000772 Published online by Cambridge University Press

APPENDIX F: FORECAST ERROR VARIANCE BEFORE AND AFTER THE GLOBAL FINANCIAL CRISIS: DECOMPOSITION OF CHANGE

	Total variances			Contribution of change in shock size				Contribution of change in impulse responses			
	(1) 1990–2006	(2) 2007–2012	(3) Change	(4) Common	(5) Spillover	(6) Domestic	(7) Total	(8) Common	(9) Spillover	(10) Domestic	(11) Total
CND	5.123***	1.238***	−3.885***	0.126	−0.033	−0.599	−0.506	−0.184	0.318	−3.513***	−3.379**
	(1.054)	(0.341)	(1.112)	(0.475)	(0.071)	(0.651)	(0.959)	(1.312)	(0.250)	(0.889)	(1.369)
FRA	4.404***	3.709***	−0.695	0.055	−0.016	0.003	0.042	−0.000	0.992	−1.728*	−0.737
	(0.758)	(1.306)	(1.504)	(0.348)	(0.247)	(0.948)	(0.935)	(0.968)	(0.824)	(1.049)	(1.585)
GER	3.442***	2.102***	−1.340	0.003	−0.002	−0.524	−0.523	−0.020	0.221	−1.018	−0.817
	(0.597)	(0.661)	(0.904)	(0.168)	(0.116)	(0.374)	(0.438)	(0.432)	(0.354)	(0.708)	(0.873)
ITA	5.766***	5.284***	−0.482	0.331	0.021	−0.990	−0.638	2.738*	0.196	−2.779**	0.156
	(1.214)	(1.819)	(2.180)	(0.662)	(0.181)	(1.362)	(1.410)	(1.546)	(0.683)	(1.418)	(2.453)
JPN	6.244***	4.239***	−2.004	0.153	−0.110	0.311	0.354	1.182	0.821	−4.361***	−2.358
	(1.274)	(1.193)	(1.737)	(0.295)	(0.183)	(0.712)	(0.775)	(0.726)	(0.665)	(1.249)	(1.726)
UK	3.132***	5.994**	2.863	0.321	−0.046	−0.025	0.249	2.201	1.294	−0.881	2.614
	(0.568)	(2.551)	(2.607)	(0.688)	(0.220)	(0.866)	(0.926)	(1.549)	(0.968)	(1.049)	(2.467)
US	9.236***	3.334***	−5.902***	1.041	−0.313	4.226***	4.954**	−8.350***	2.812***	−5.318***	−10.856***
	(1.687)	(1.059)	(1.998)	(0.888)	(0.306)	(1.584)	(2.015)	(2.968)	(0.989)	(1.716)	(2.967)

Note: Columns (1) and (2) show the 12-months-ahead forecast error variance decomposition of inflation uncertainty for two subsamples and the difference between the two subsamples based on an FSVAR estimation with one common factor, four autoregressive lags, and one lag of the remaining six countries in each equation. Changes are reported in column (3). Columns (4) to (6) report the change in the magnitude of the common shock, country-specific shocks from abroad (spillover), and domestic shocks. Column (7) shows the total contribution of the change in shock size. Columns (8) to (10) report contributions of change in propagation of common shocks, country-specific shocks from abroad (spillover), and domestic shocks in the domestic economies. The sum of these contributions is reported in column (11). The values are multiplied by 100, and bootstrap standard errors are reported in parentheses.