Monetary Policy Uncertainty and Unionized Labor Markets:

Empirical Evidence

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Abstract

Recent theoretical research has studied extensively the link between wage setting and monetary policymaking in unionized economies. This paper addresses the question of the role of monetary uncertainty from an empirical point of view. Our analysis is based on a simple model that derives the influence of monetary uncertainty on wage setting. We test our model with data for the G5 countries. A central finding is that monetary policy uncertainty has a negative impact on wage growth in countries where wage setting is relatively centralized. This is consistent with recent theoretical approaches to central bank transparency and wage setting.

Keywords: Monetary policy uncertainty, centralized wage setting, union behavior.

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1. Introduction

The role of transparency of monetary policy decisions has recently received considerable attention in the theoretical literature. Most of the literature was concerned with the ability of financial markets to correctly anticipate moves of the monetary authority. The main argument is that it increases the effectiveness of monetary policy as it improves the market participants understanding of the intentions of the central bank. It also increases the credibility of monetary policy and makes it easier for the central bank to communicate its decisions and to explain changes in its behavior or why it has failed to meet its goals (see Blinder et al. 2001, Geraats 2002). On the other hand, Sorensen (1991) has argued that uncertainty may have some positive effects when it affects the behavior of other macroeconomic players. According to Sorensen's argument monetary uncertainty may lead to wage restraint and hence to lower inflation and unemployment. Recently, Grüner (2002) and Grüner and Hefeker (2002) have also focused on the influence of uncertainty about the central bank's behavior on macroeconomic outcomes. If labor unions can not be certain how their wage setting behavior will impact the central banks behavior, they tend to be less aggressive and more cautious in formulating wage demands.

The present paper builds on the second approach and addresses the question of the role of monetary uncertainty primarily from an empirical point of view. We try to answer the question whether central banks should be concerned with establishing the best possible predictability of their behavior. We test the theory of wage restraints in a situation of high monetary uncertainty empirically using data from the G5 countries (US, France, Japan, the UK and Germany) and analyze the influence of uncertainty about the central bank's behavior on wage inflation in these economies.

We find that higher monetary policy uncertainty has a negative effect on wage inflation in the continental European economies and in Japan. We are unable to detect such a disciplinary effect in the UK and the US. One possible explanation for this, supported by the theoretical set up that we develop in Section 2 of the paper, is the different degree of wage coordination in the economies considered. While Germany, France and Japan can be considered as relatively coordinated, this is not the case for UK and US, which are characterized by more atomistic labor markets (Hargreaves Heap 1994, OECD 1994).

The paper is related to recent theoretical research work on the link between wage setting and monetary policymaking in unionized economies. Focusing on the characteristics of the central bank, Skott (1997), Sorensen (1991), Grüner and Hefeker (1999), Guzzo and Velasco (1999), and Cukierman and Lippi (1999, 2001) have studied the impact of monetary policy and labor market institutions on macroeconomic outcomes. Here the main issue has been whether a conservative, i.e. highly inflation averse, central bank should be preferred over a liberal, i.e. employment concerned, central bank. These papers invalidate the standard Rogoff (1985) result that a conservative central banker is welfare increasing for the economy when labor markets are decentralized. Considering the case of non-atomistic labor unions, a case can be made that liberal central bankers might discipline inflation averse labor unions because unions fear too strong an inflationary response by the liberal central bank.[1]

The present paper is also linked to the literature on the desirability of central bank transparency. Cukierman and Meltzer (1986) have made the case for so called „creative ambiguity“, arguing that only unanticipated monetary policy can be effective. Given that rational individuals anticipate the reaction of the central bank, monetary policy becomes powerless except in reaction to shocks. Focusing on the role of asymmetric information, Cukierman (2000) and Jensen (2002) have argued that the desire of the central bank to be transparent can be counterproductive. By sending too many signals (bits of information) to the public and the financial markets, they may create excessive volatility in the financial markets. In order to avoid these strong movements, the central bank may be forced into inactive behavior.

To our knowledge, there exist only few empirical approaches to the question of how the central bank characteristics influence the wage setting behavior of unions. Grüner (1995) is one contribution that takes into account that both the monetary authorities and the trade union characteristics affect wages and prices. More recently, Cukierman and Lippi (1999) have analyzed how the degree of decentralization of wage setting and central bank independence affect inflation and unemployment. They find that at low levels of central bank independence, the hump shaped relation between unemployment and wage centralization postulated by Calmfors and Driffill (1988) can be confirmed but that this relation disappears at higher levels of central bank independence.

Chortareas et al. (2001) analyze empirically a specific type of transparency, namely the publication of inflation forecasts and other forward-looking indicators. Performing an analysis in a cross-section of countries, where they relate the details of central banks' forecasts with the inflation rate, they find that, after controlling for a number of institutional characteristics, an increase in the forecast detail is associated with lower average inflation.

This paper instead focuses on the question how monetary uncertainty affects wage growth. Empirically, we work along the time dimension, i.e. we let monetary uncertainty vary over time, and compare our results across the G5 countries.

2. A Simple Model of Monetary Policy Uncertainty and Wage Setting

A simple model, developed by Grüner and Hefeker (2002), analyzes the relation between monetary policy uncertainty and wage setting by labor unions. Consider an economy which is populated by n unions that interact with the monetary authority. All unions simultaneously fix their nominal wage demands, taking the expected rate of inflation into account. After this, the monetary authority sets its policy and employment is determined according to labor demand.

[1] The important assumption is the inflation aversion of labor unions. For a micro foundation of

Let the demand for labor in sector i be given as

$$L_i^d = \frac{L}{n}\left(1-\left(w_i - p\right)\right), \tag{1}$$

where L is the total labor supply in the economy.[2] Total labor demand is given as

$$L^d = \sum_{i=1}^{n}\frac{L}{n}\left(1-\left(w_i - p\right)\right). \tag{2}$$

Thus, the unemployment rate follows as

$$u = \frac{L - L^d}{L} = \frac{L - \sum_{i=1}^{n}\frac{L}{n}\left(1-\left(w_i - \pi\right)\right)}{L} = \frac{L - L + L(\overline{w} - \pi)}{L} = \overline{w} - \pi. \tag{3}$$

where $\overline{w} = \frac{1}{n}\sum_{i=1}^{n} w_i$. The rate of unemployment in sector i is

$$u_i = w_i - p\ . \tag{4}$$

The objectives of the labor union in sector i are given as

$$U_i = w_i - p - \frac{A}{2}u_i^2\ . \tag{5}$$

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this inflation aversion, see Berger et al. (2002).

Hence, we assume that labor unions are concerned with maximizing the real wage of their members in the respective sector. They are also averse to unemployment in their sectors. The relative weight unions place on the employment aim is $\frac{A}{2}$, which is assumed to be the same for all labor unions.

The central bank's objective function in turn is given as

$$C = -\left[Ip^2 + u^2\right]. \tag{6}$$

The central bank aims at holding inflation at zero and avoiding unemployment. The inflation aversion of the central bank is measured as I. Inserting (3) in (6) and optimizing with respect to its policy variable, which for simplicity is taken to be the rate of inflation, its optimal policy follows as

$$p = \frac{\overline{w}}{1+I} \equiv b\overline{w}. \tag{7}$$

The central wage reacts to the average wage demands in the economy with an increase in the rate of inflation, $b \leq 1$. It is instructive to look at some special cases of I: For I = 0, b will be unity, i.e. there is a full pass-through of wage setting to the inflation rate. For I = 1, b will be one-half. Finally, for I = ∞, b equals zero, i.e. wages have no impact on inflation.

This reaction of the central bank to the behavior of the labor unions is what we consider to be uncertain for the labor union. The reaction of the central bank is uncertain if the central bank is not transparent in terms of its objective function and behavior. Hence, if a central bank is

[2] Inflation and price level in period t are the same when normalising P_{t-1} appropriately.

nontransparent, unions form expectations about the banks reaction to their wage setting. This is given as $E(b)=\hat{b}$ and $Var(b)=\sigma_b^2$.

Taking these properties of the expected central bank behavior into account, the unions objective function can be rewritten, in expected values, as

$$E(U_i)=E\left[w_i-b\overline{w}-\frac{A}{2}\left(w_i-b\overline{w}\right)^2\right]. \tag{8}$$

This leads to

$$\frac{\partial E(U_i)}{\partial w_i}=1-\frac{\hat{b}}{n}-\frac{A}{2}\left[2w_i+2\left(\sigma_b^2+\hat{b}^2\right)\overline{w}\cdot\frac{1}{n}-2\hat{b}\left(\overline{w}+w_i\cdot\frac{1}{n}\right)\right]=0, \tag{9}$$

where we have made use of the fact that $\mathrm{E}\left(\mathrm{b}^2\right)=\sigma_\mathrm{b}^2+\hat{\mathrm{b}}^2$.

In a symmetric equilibrium ($w_i = \overline{w} = w$) we find that each labor union will set a wage demand of

$$w=\frac{n-\hat{b}}{A\left(\left(n-\hat{b}\right)\left(1-\hat{b}\right)+\sigma_b^2\right)}. \tag{10}$$

The equilibrium wage in (10) is greater than (or in the case of one union, equal to) zero, since it follows from $\hat{\mathrm{b}}\leq 1$ that n > $\hat{\mathrm{b}}$. Several conclusions can be derived from this equation. First, as Sorensen (1991) has already established, a decrease in the transparency of the central bank, which is reflected in an increase in σ_b^2 lowers the wage demands of the labor unions since

$$\frac{\partial w}{\partial \sigma_b^2}=\frac{-\left(n-\hat{b}\right)A}{\left[A\left(\left(n-\hat{b}\right)\left(1-\hat{b}\right)+\sigma_b^2\right)\right]^2}<0. \tag{11}$$

The intuition for this is that more uncertainty makes the labor union more cautious in its wage demands. If it can not be certain how an increase in wages will translate into monetary policy and thus, via labor demand, into employment in its sector, the union will be more reluctant to demand higher wages.

Second, an increase in the number of labor unions in the economy has a positive effect on wage demands because

$$\frac{\partial w}{\partial n} = \frac{A s_b^2}{\left[A\left(\left(n-\hat{b}\right)\left(1-\hat{b}\right)+s_b^2\right)\right]^2} > 0 . \tag{12}$$

Here, the intuition is that with more decentralized wage setting unions internalize less the influence of their own behavior on the central bank's reaction. It is clear that this effect vanishes in a non-stochastic environment where unions can perfectly predict the behavior of the central bank. Then the characteristics of the bank play no role since the unions can always adapt their wage demands to the central bank's reaction. If, however, the reaction of the bank is stochastic it is important how many unions there are. If there is only one union, this union will internalize the uncertain reaction of the central bank in its wage setting. But this internalization effect vanishes as the number of unions increase.

Third, one can explore the effect of changes in the number of labor unions on the influence of uncertainty of wage demands. This is given as

$$\frac{\partial^2 w}{\partial s_b^2 \partial n} = \frac{-A^3\left[2\left(n-\hat{b}\right)\left(1-\hat{b}\right)+\left(\left(n-\hat{b}\right)\left(1-\hat{b}\right)+s_b^2\right)^2\right]}{\left[A\left(\left(n-\hat{b}\right)\left(1-\hat{b}\right)+s_b^2\right)\right]^4} < 0 , \tag{13}$$

which states that the moderating influence of uncertainty about the central bank's reaction on wage demands is weakened by an increase in the number of labor unions. This is not surprising, given that the two effects have opposite influences on wage demands when considered separately.

Thus, the model predicts that the uncertainty of central bank reaction to wage demands has a dampening effect on wage demands but this dampening effect will be less if there is an atomistic labor market. These are the two hypotheses that we are going to test in the next sections of our paper.

Our empirical analysis uses the variance of the expected interest rate as a proxy for monetary uncertainty. Our theoretical model instead uses the variance of the reaction parameter of the central bank σ_b^2. Thus our model and the empirical implementation are only compatible if there is a positive correspondence between the variance of the policy instrument and the variance of central bank behavior. This is the case if the derivative of the variance of inflation (our monetary policy variable) with respect to the variance of the preference parameter is positive. As has been pointed out by Grüner (2002) this is condition is fulfilled if σ_b^2 is sufficiently small. To see this, consider that

$$\frac{\partial \sigma^2}{\partial \sigma_b^2} = \frac{\left(n - \hat{b}\right)}{\left[A\left(\left(n - \hat{b}\right)\left(1 - \hat{b}\right) + \sigma_b^2\right)\right]^2} - 2\sigma_b^2 \frac{A\left(n - \hat{b}\right)}{\left[A\left(\left(n - \hat{b}\right)\left(1 - \hat{b}\right) + \sigma_b^2\right)\right]^3}, \tag{14}$$

which is positive for $\left(n - \hat{b}\right)\left(1 - \hat{b}\right) > \sigma_b^2$.

Hence, for each value of $\hat{b}$ there is an upper bound on the variance such that the relation between preference uncertainty and monetary policy uncertainty is positive.

3. Constructing an Indicator of Monetary Policy Uncertainty

In the first step of the empirical analysis we have to construct an indicator for monetary policy uncertainty. We assume that the stance of monetary policy can be proxied by a short-term interest rate. Using the expectations theory of the term structure of interest rates, we include long-term interest rates in the explanation of short-term rates (see Balduzzi et al. (1997), Gerlach and Smets (1997), Hsu and Kugler, (1997), Nautz (2000)). Unit-root testing reveals that both short- and long-term interest rates are I(1) variables (results omitted). We estimate the relationship between these variables in a VECM (vector error correction model). The estimation period employs monthly data from 1979:1 to 1998:12. It was chosen to coincide with the establishment of the European Monetary System in Europe until the start of European Monetary Union. Details about the data sources are given in the Appendix.

The results of the cointegration analysis for these two variables using the reduced-rank method by Johansen (1988) are summarized in Table 1.

Table 1: Cointegration analysis of interest rates (1979:1 to 1998:12)

	France	Germany	Japan	UK	US
Lags	6	6	6	4	6
Eigenvalues	0.07, 0.001	0.05, 0.009	0.09, 0.002	0.06, 0.001	0.05, 0.02
Trace/Max	17.8*, 17.6*	14.4(*), 12.4(*)	22.9**, 22.3**	16.1*, 16.0*	13.9(*), 10.4
β	1, -1.08	1, -1.00	1, -1.14	1, -1.03	1, -0.99
α	-0.12, -0.003	-0.05, -0.003	-0.11, 0.01	-0.12, -0.005	-0.06, 0.01
$\beta = (1, -1)$ & $\alpha = (u, 0)$	$Chi^2(2)=0.72$	$Chi^2(2)=0.07$	$Chi^2(2)=3.1$	$Chi^2(2)=0.30$	$Chi^2(2)=1.10$

Notes: Only the first eigenvalue is significant. An unrestricted constant is always included in the model. The equations for Germany contain impulse dummies for 1981:2 and 1981:3 and for France for 1992:9 as unrestricted variables. In the case of the UK, the estimation period has to start in 1978:1 to get significant cointegration results. Critical values are taken from Osterwald-Lenum (1992). **, *, (*) indicate significance at a 1%, 5% and 10% level, respectively.

The table has the results for each country ordered by columns. The second line reports the number of lags needed to avoid autocorrelation in the residuals. Next, the two eigenvalues are provided. Since the smaller eigenvalue is never significant, we report trace and maximum eigenvalue statistics for the larger one only. We cannot reject the hypothesis of one cointegration vector based on a 10% significance level. The estimated cointegration vectors, denoted by β, and the adjustment vectors, denoted by α, are given in the next lines. Finally, the last line contains the result of testing restrictions on β and α. In particular, it is tested whether the coefficient linking short-term and long-term interest rates is unity, and whether the long-term interest rate is weakly exogenous (see Johansen, 1992). In none of the cases do we have to reject the hypothesis. This implies that we can compute the following error-correction term:

$$ECM_t = \text{Short-run interest rate}_t - \text{Long-term interest rate}_t \quad (15)$$

Moreover, tests on the adjustment vector reveal (results omitted) that the long-term rate is weakly exogenous for all countries.[3] This implies that we can continue modeling within a single-equation context.

Next, we estimate an error-correction model for each of the G5 countries. Diagnostic testing reveals that residuals are showing significant ARCH (autoregressive conditional heteroscedasticity) effects. Therefore, we augment the error correction model by a GARCH(1,1) framework (Bollerslev, 1986) that takes account of the observed ARCH effect. Since we are not really interested in the actual estimation results, Table 2 reports important diagnostic information. The first line of the table summarizes the results of the ARCH test. It is apparent that the GARCH(1,1) model is successful in removing the ARCH component from the equation. In addition, no significant autocorrelation can be found in the residuals (AC test). There is, however, non-normality in the residuals and if we were to undertake any inference in the model we would need to rely on robust standard errors (see Bollerslev and Wooldridge, 1992).

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[3] All omitted results are available upon request.

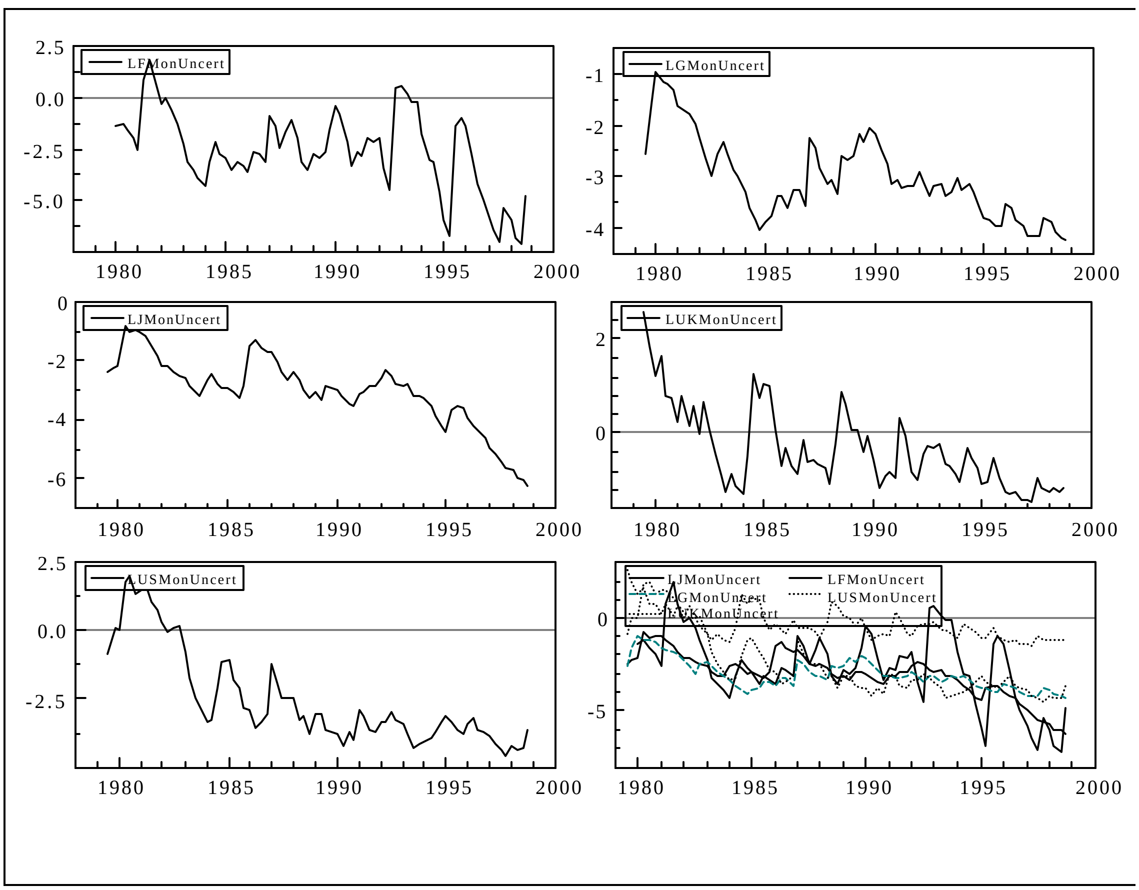
Table 2: Diagnostic testing of error-correction GARCH(1,1) model

	France	Germany	Japan	UK	US
ARCH	F(2,213) =0.02	F(2,213)=0.5	F(2,215)=0.8	F(2,221)=0.6	F(2,214)=1.0
AC	$Chi^2(31)=40$	$Chi^2(31)=20$	$Chi^2(31)=41$	$Chi^2(33)=39$	$Chi^2(31)=31$
Normality	$Chi^2(2)=3342$**	$Chi^2(2)=29$**	$Chi^2(2)=59$**	$Chi^2(2)=53$**	$Chi^2(2)=41$**

Notes: ARCH is the Engle-test for ARCH effects. AC is a Portmanteau test for autocorrelation. Normality is the Doornik-Hansen-test for normally distributed residuals. In the case of France, Japan, and US, we had to impose a restriction on the GARCH coefficients to ensure stationarity. **, *, (*) indicate significance at a 1%, 5% and 10% level, respectively.

The interesting part of this model is (the logarithm of) the estimated conditional variance for the short-term interest rate, which is taken as our indicator for monetary policy uncertainty (see Sauer and Bohara (1995) for a different way to model monetary uncertainty).

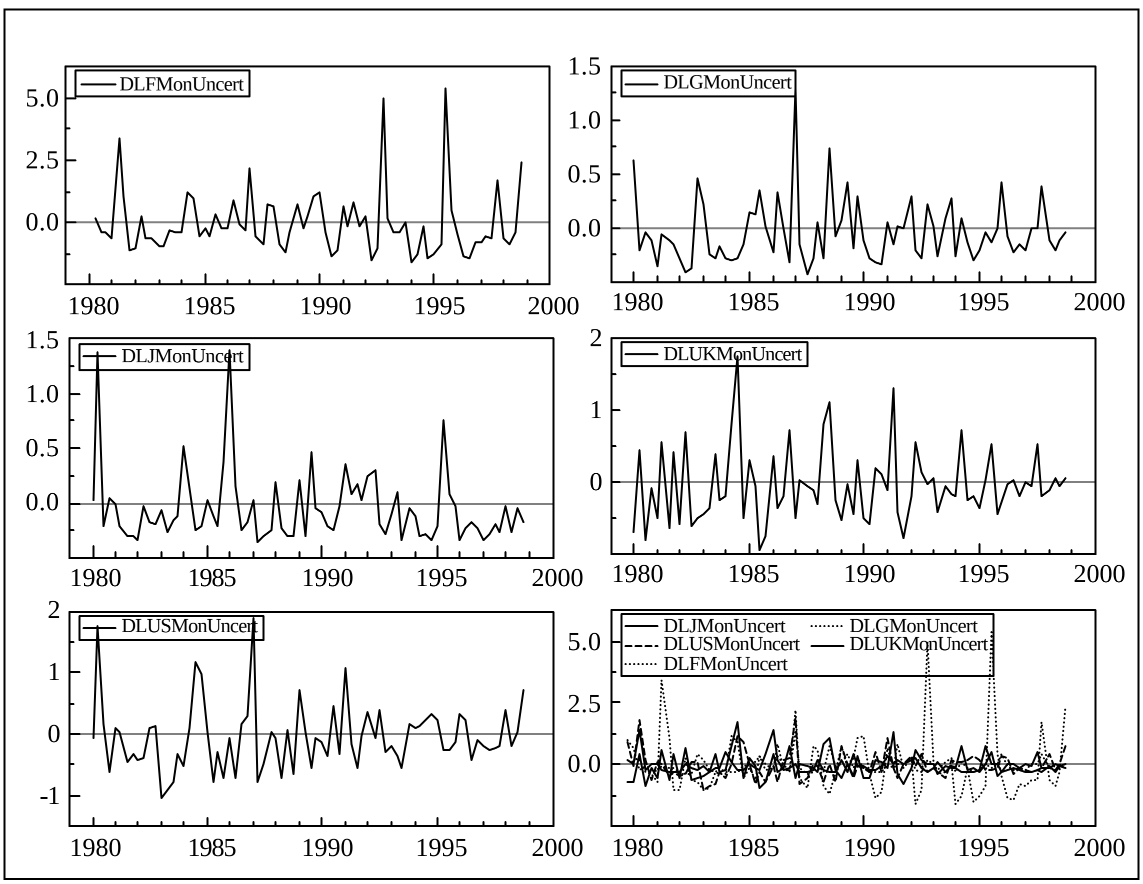
Figure 1: Conditional variance of short-term interest rate (in logs)



From Figure 1 it is apparent that the conditional variance decreases over the course of the sample period. Using conventional unit-root tests (results omitted), it can be shown that this indicator for monetary uncertainty exhibits stochastic non-stationary, or, to be more specific, it seems to be I(1). That the processes were on or close to the unit circle was already apparent from the estimation of the GARCH processes. Since the log of the conditional variance shows the properties of an I(1) process, and this causes problems for inference, it is prudent to map it into I(0)-space by computing first differences. *Thus, we measure monetary policy uncertainty as the rate of change of the conditional variance*.

Figure 2 plots this measure of monetary uncertainty over time. It is apparent that the series are now stationary.

Figure 2: Monetary uncertainty indicators (change of (log) conditional variance)



Comparing the indicators for monetary uncertainty shows that fluctuations generally decline over time with the exception of France. As the last graph shows, the series for France displays large fluctuations even at the end of the sample period.

Table 3 provides some descriptive statistics of the indicator for monetary policy uncertainty. We find that the means of the series have a negative sign and are relatively similar in size. This implies that monetary uncertainty is decreasing throughout the countries in our sample, with the decline in Japan being the strongest and with the UK at the opposite end. Reflecting the behavior of the series in Figure 2, the standard deviation of France is the highest. The UK and the US are intermediate cases, while Germany and Japan show the lowest variations.

Table 3: Means, standard deviations, and correlation coefficients for Monetary Uncertainty Indicators

	France	Germany	Japan	UK	US
Means	-0.045	-0.044	-0.053	-0.032	-0.049
Standard deviations	1.284	0.284	0.333	0.502	0.540
Correlation coefficients					
France	1				
Germany	0.27	1			
Japan	0.08	-0.06	1		
UK	0.05	-0.16	0.05	1	
US	0.15	0.23	0.20	0.13	1

The correlation between France and Germany is the strongest. Arguably, this reflects the interest rate linkage among these countries within the EMS. Most other countries do not show much correspondence, except for a negative correlation between UK and Germany. Noteworthy, however, is the relatively high correlation coefficient of the US with all other countries in the sample. This underlines the importance of the US with regard to interest rate setting.

4. Modeling nominal wage growth

Next, we test the influence of our monetary policy indicator in a regression explaining nominal wage growth. Since wages are only available as quarterly series, we shift the analysis to this frequency. The dependent variable wage inflation is defined as the first difference of the log of the wage index. Although some of the series still seem to show some trending behavior, they are difference stationary according to standard unit root tests (results omitted).

The highest average wage inflation can be observed for the UK, which also displays the largest standard deviation. However, absolute differences between countries tend to be relatively small. At the end of the sample, we find wage inflation being relatively strong in the US and the UK, and relatively small in Japan. This reflects the different output growth experience of these countries over this time period.

As a modeling vehicle for wage inflation, we utilize the Phillips-curve framework. Moreover, the actual modeling uses the general-to-specific approach (Hendry, 1990). The general model consists of four lags of nominal wage growth, deviation of actual unemployment from NAIRU (non-accelerating inflation rate unemployment), and the monetary uncertainty indicator. The NAIRU is time varying and is constructed by subtracting the trend in unemployment, as measured by a Hodrick-Prescott filter, from the actual unemployment rate series (see Estrada et al. 2000 and Staiger et al. 2001 for recent attempts to estimate the NAIRU). We expect that unemployment reduces wage growth if it is above the NAIRU.

We want to make sure that the estimated models are reasonably stable over time. Therefore, we reserve the last eight observations for out-of-sample stability tests. Starting with the general model, more parsimonious specifications were derived in a consistent testing down-process at a (nominal) 5% significance level.

Table 4 contains the results of diagnostic tests, showing that the simplified models are valid reductions, as they are still congruent representations of the data generating process. The only rejection of a null hypothesis occurs in the case of Japan in the out-of-sample Chow-test statistics. Here the main culprit appears to be the deflationary economic environment at that time in Japan, which apparently causes the predictive failure.

Table 4: Diagnostic statistics of nominal wage growth models

	France	Germany	Japan	UK	US
AC (1-5)	F(5,43) = 2.30	F(5,56) = 1.25	F(5,56) = 1.13	F(5,48) = 2.16	F(5,55) = 0.74
ARCH(1-4)	F(4,40) = 0.90	F(4,53) = 0.25	F(4,53) = 0.74	F(4,45) = 0.59	F(4,52) = 0.44
Normality	$Chi^2(2)$ = 0.15	$Chi^2(2)$ = 2.46	$Chi^2(2)$ = 4.50	$Chi^2(2)$ = 3.11	$Chi^2(2)$ = 0.81
Hetero	F(14,33) = 0.71	F(6,54) = 0.65	F(6,54) = 1.38	F(4,48) = 1.55	F(7,52) = 0.15
Chow	F(8,48) = 0.58	F(8,61) = 1.87	F(8,61) = 3.05**	F(8,53) = 0.59	F(8,60) = 0.42
Instability uncertainty	0.41	0.04	0.26	n.a.	n.a.

Notes: AC is an LM test for autocorrelation from lag 1 to 5. ARCH is the Engle-test for ARCH effects on lags 1 to 4. Normality is the Doornik-Hansen-test for normally distributed residuals, Hetero is the White-test for heteroscedasticity (using squared values only). Chow is a Chow-test on the 8 out-of-sample observations, instability uncertainty is a Hansen-test for stability of the indicator for monetary uncertainty. **, *, (*) indicate significance at a 1%, 5% and 10% level, respectively.

Table 5 contains the estimated results of the simplified models. Most models display only few significant effects. Usually, we find a negative impact of the deviation of the actual unemployment rate from the NAIRU and a positive coefficient on the lagged endogenous variable. The time pattern of these effects varies from country to country, probably reflecting different labor market institutions and adjustment dynamics. In the case of France, we cannot reject the hypothesis that the sum of the coefficients on the NAIRU variable is equal to zero (F(1,48) = 1.39).

Table 5: Modeling nominal wage growth 1979:1 to 1998:4

	France	Germany	Japan	UK	US
$\Delta Wage_{t-1}$		-0.257*		0.497**	0.572**
		(0.122)		(0.119)	(0.090)
$\Delta Wage_{t-2}$	0.089		0.447**		
	(0.082)		(0.105)		
$\Delta Wage_{t-3}$	0.339**				
	(0.080)				
$\Delta Wage_{t-4}$	0.290**				
	(0.070)				
$NAIRU_t$			-0.004(*)	-0.002(*)	-0.0006
			(0.002)	(0.001)	(0.0008)
$NAIRU_{t-1}$		-0.006**			
		(0.001)			
$NAIRU_{t-2}$	-0.022**				
	(0.005)				
$NAIRU_{t-3}$	0.031**				
	(0.009)				
$NAIRU_{t-4}$	-0.011*				
	(0.005)				
$Uncertainty_t$	-0.0009*				-0.0009
	(0.0004)				(0.0011)
$Uncertainty_{t-1}$					
$Uncertainty_{t-2}$			-0.0023*		
			(0.0010)		
$Uncertainty_{t-3}$					
$Uncertainty_{t-4}$		-0.003(*)			
		(0.002)			
Constant	0.002*	0.013**	0.004**	0.008**	0.006**
	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)
SE	0.0035	0.0044	0.0028	0.0055	0.0043
R^2	0.77	0.30	0.30	0.38	0.58
F-test	F(7,48) = 22.7**	F(3,61) = 8.88**	F(3,61) = 8.70**	F(2,53) = 16.1**	F(4,60) = 20.7**
Cases	56	65	65	56	65
Test of model reduction	F(7,41) = 1.01	F(11,50) = 0.89	F(11,50) = 0.80	F(12,41) = 1.53	F(11,49) = 1.58

Notes: Wage data for France and the UK start in 1983. **, *, (*) indicate significance at a 1%, 5% and 10% level, respectively.

Regarding the influence of monetary uncertainty, we find two groups of countries: France, Germany and Japan show a negative influence of the change of monetary uncertainty on wage setting, while we do not find such an effect in the UK or US. [4]

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[4] The marginal significance level for the German monetary uncertainty variable is 0.053.

Arguably, France, Germany, and Japan are rather characterized by coordinated wage setting, while in the US and the UK decentralized and uncoordinated wage setting is more prominent (OECD 1994, Hargreaves Heap 1994). Hence, the results confirm our theoretical prior above, namely that for those countries where trade unions presumably take spill-over effects of wage decisions into account, monetary uncertainty reduces wage inflation.

To ensure that this conclusion is robust, we conduct a number of tests:

Firstly, we use an insample stability test to ensure that the effect of monetary uncertainty on wage growth is reasonably stable. The test by Hansen (1992) allows to test specific coefficients in a model for stability. As is apparent from the last line of Table 4, none of the test statistics are significant at a 5% level.

Secondly, it may be argued that our wage growth equations do not adequately account for the effect of expected price changes and for the influence of productivity. Therefore, we compute omitted variable tests for contemporaneous values of these variables as a robustness check. We test three indicators for future inflation: First, a contemporaneous inflation rate derived from an instrumental variable regression with four lags of inflation and the contemporaneous value plus four lags of the output gap. Second, the contemporaneous output gap. Third, the current change in labor productivity to account for the variation in productivity.

In the case of France, the instrumented current inflation rate is significant at a 10% level. It shows a positive sign when including it in the model, while leaving all other variables basically unaffected. In the German model, none of these variables is significant. For Japan we find a significant effect (at the 5% level) of labor productivity. Including this variable in the model results in a positively signed coefficient. It reduces the p-value for monetary uncertainty to 0.053 but cannot improve on the stability properties of the model as tested by the out-of-sample Chow-test. In the UK labor productivity is significant at the 10% level, and it enters the equation with a positive sign. Since monetary uncertainty is not significant anyway, this does not change the conclusion. Finally, for the US we get a significant impact at the 10%-level for the inflation variable, although including it would result in a negative sign of the coefficient. More relevant is the change in labor productivity, which is significant at a 1% level. It enters with the right sign but leaves both the coefficients on monetary uncertainty and the NAIRU insignificant.

Thirdly, some researchers believe that it is useful to use White's or Andrews' robust standard errors (White, 1980, Andrews, 1991) as a precaution against invalid inference, even though no obvious heteroscedasticity and/or autocorrelation has been detected. In our case, using these robust standard errors would lead, in all instances, to more significant estimates of the parameter on monetary uncertainty, without changing the general conclusion for the sample of G5 countries. To summarize, the results of our models hold up quite well against these tests of robustness.

5. Conclusion

Several recent theoretical papers explored the link between wage setting and monetary policymaking in unionized economies. This literature states that monetary uncertainty may have a disciplinary impact on wage growth in countries where labor unions internalize the influence of their behavior on the monetary policy of the central bank. We tested the model for five countries, Germany, France, Italy, the UK and the US. Our empirical analysis is consistent with the view that monetary policy uncertainty may lead to wage restraint.

This has important consequences both for the design of monetary policy institutions and for the strategy of monetary policy. When wage setting is not too decentralized, monetary policy may increase employment by using "creative ambiguity" as argued by Cukierman and Meltzer (1986). Unions wage demands are disciplined if they internalize the central bank's reaction to their wage demands. Our model also implies, however, that if trade unions are already very conscious about employment, "creative ambiguity" in monetary policy is not going to reduce unemployment very much.

We consider the empirical results as preliminary that should be followed by more extensive research. Further empirical research could study the robustness of this result using alternative empirical indicators for uncertainty (see, e.g., Sauer and Bohara 1995). One such possibility would be to use measures of how futures rates precede changes of the monetary policy stance as an indicator of monetary transparency (see Blinder et al. 2001). Moreover, it would be useful to disentangle uncertainty that arises from monetary policy itself and uncertainty that arises from the rest of the economic environment. Finally, the degree of conservativeness of monetary policy may be associated with the size of nominal wage claims. It would be useful to generate time varying estimates of these factors and to use them in a more encompasssing empirical analysis.

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Appendix

Data description

Interest rate data (monthly series):

IMF International Financial Statistics (Short-term rate: …60b, long-term rate: …61).

Unemployment data (quarterly series):

Datastream

Wage data (quarterly series):

France: Datastream Hourly Wage Rate, All Activities, SA, in FF.

Germany: Statistisches Bundesamt (Hourly Wage Earnings, West-Germany, 1995=100, SA using Census X-12).

Japan: IMF International Financial Statistics (Wages: Monthly earnings), SA using Census X-12).

UK: IMF International Financial Statistics (AV Earn Prod Ind SA, 11265..CZF…).

US: Datastream US Wages & Salaries (AR) CURA in US$, SA using Census X-12.