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by Hans Peter Grüner [2], Bernd Hayo [3] and Carsten Hefeker [4]

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CONTENTS

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 both an empirical and theoretical point of view. Our analysis is based on a simple model that derives the influence of monetary uncertainty on unionized wage setting. We construct an indicator of monetary policy uncertainty and test our model with data for the G5 countries. The central finding is that monetary policy uncertainty has a negative impact on nominal wage growth in countries where wage setting is relatively centralized. This result is consistent with recent theoretical approaches to central bank transparency and wage setting.

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

JEL-Classification: E58

Non-technical Summary

Can ambiguous monetary policy create jobs? Recent theoretical work has argued that monetary uncertainty may have some positive effects when it influences the behavior of other macroeconomic players. Accordingly, monetary uncertainty may lead to wage restraint and hence to lower inflation and unemployment. If labor unions can not be certain how their wage setting behavior will affect the central bank's behavior, they tend to be less aggressive and more cautious in formulating wage demands. According to this argument, risk averse labor unions take into account that increased wage demands could lead to a higher variance of inflation and employment when the central bank's reaction is less predictable. Therefore, ambiguous monetary policy reduces wage inflation if wage setting is coordinated. This theoretical argument has so far not been scrutinized empirically.

In the present paper it is analyzed whether and when wage effects of monetary ambiguity may exist. The empirical analysis proceeds in two steps. First, we derive an indicator of monetary policy uncertainty based on the conditional variance of an equation explaining the short-term interest. Second, we use this indicator of monetary policy uncertainty and test its influence on nominal wage inflation in the G-5 industrial economies (US, France, Japan, the UK and Germany).

It is found that higher monetary policy uncertainty has a significantly negative effect on wage inflation in the continental European economies and in Japan. No disciplinary effect was detected in the data on the UK and the US. Based on the theoretical model, we argue that this is due to a lower degree of wage coordination in the latter countries. In other words, highly centralized unions realize that their wage demands affect the overall wage inflation and therefore monetary policy more strongly. Thus, the degree of monetary policy uncertainty should have a stronger impact on more centralized labor markets. 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.

The results presented in the paper stand in contrast to the recommendations derived in another part of the literature on the optimal degree of transparency in monetary policy making. Most of the literature is concerned with the ability of financial markets to correctly anticipate moves of the monetary authority. The main argument is that transparency 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. In terms of policy recommendations, the present paper does not postulate that monetary policy making ought to be as ambiguous as possible. But it provides a theoretical argument and some empirical evidence that it may also be suboptimal to be as transparent as possible. In other words, we would like to provide a word of caution to those monetary policy commentators that argue in favor of utmost transparency. After all, there may still be a role for "creative ambiguity" in monetary policy making.

1. Introduction

Can ambiguous monetary policy create jobs? A recent theoretical literature argues that more monetary uncertainty reduces trade unions' incentives to claim higher wages (Sorensen, 1991, Grüner, 2002a, 2002b). According to this theoretical argument, risk averse labor unions internalize that increased wage demands could lead to a higher variance of inflation and employment when the central bank's reaction is less predictable. Therefore, ambiguous monetary policy reduces wage inflation. This theoretical argument has, however, not been scrutinized empirically so far. In the present paper we use a simple estimation procedure in order to analyze whether and when wage effects of monetary ambiguity may exist. We construct an indicator of monetary policy uncertainty and test its influence on wage inflation in the G-5 industrial economies (US, France, Japan, the UK and Germany).

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. More centralized unions internalize that their wage demands influence monetary policy more strongly, and therefore the degree of monetary policy uncertainty should have a stronger impact on more centralized labor markets. 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 role of the degree of transparency in monetary policy decisions has recently received considerable attention in the theoretical literature. Most of the literature is 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).

Cukierman and Meltzer (1986) instead 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.

Likewise, Sorensen (1991) has argued that uncertainty may have some positive effects when it influences 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 (2002a) has 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 influence the behavior of the central bank, they tend to be less aggressive and more cautious in formulating wage demands. We build on the second approach and address the question of the role of monetary uncertainty from both a theoretical and an empirical point of view.

The paper is also 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 partly 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 highly centralized wage setting, 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]

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 authority's and the trade union's 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.

The plan of the paper is as follows: In the next section, we develop the theoretical framework for studying the effect of monetary policy uncertainty on wage setting. Section 3 discusses the construction of an empirical indicator for monetary policy uncertainty. The impact of this uncertainty on wage inflation is analyzed in Section 4. The last section summarizes the main findings and discusses their policy implications.

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[1] The important assumption is the inflation aversion of labor unions. For a micro foundation of this inflation aversion, see Berger et al. (2004).

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

Building on Sorensen (1991), we present a simple model of the interaction between various sectoral trade unions and a single central bank which helps to understand the impact of monetary uncertainty on wage setting under different degrees of trade union centralization. We 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 a given labor demand schedule.

Let the demand for labor in sector i be given as

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

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

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

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

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

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

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

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

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

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

The central bank aims to hold inflation at zero and avoid 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

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

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 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

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

This leads to

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

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

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

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

The equilibrium wage in (10) is greater than (or in the case of one union, equal to) zero, since it follows from $\hat{b} \leq 1$ that $n > \hat{b}$. 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

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

The intuition for this is that greater 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

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

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. When 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

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

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 because more centralized unions internalize the impact of their wage demands on monetary policy, and therefore the degree of monetary policy uncertainty has a larger impact on more centralized labor markets. Thus, the model predicts that the uncertainty of central bank reaction to wages 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 the 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 (2002a) 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 uncertainty about the central bank's reaction parameter and monetary policy uncertainty is positive.

3. Constructing an Indicator of Monetary Policy Uncertainty

The first step of the empirical analysis is the construction of an indicator for monetary policy uncertainty. We assume that the stance of monetary policy can be proxied by a short-term interest rate (see Borio, 1997). As a framework for explaining short-term interest rates, we use the expectations theory of the term structure of interest rates, i.e. 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) using monthly data from 1979:1 to 1998:12. Thus, the sample starts in the year when the ERM was formed and stops with the creation of EMU. This choice of the sample excludes major breaks in the monetary regime for these countries that may bias the results. Details about data sources are given in the Appendix.

The results of the cointegration analysis for the two interest rate variables using the reduced-rank method proposed by Johansen (1988) are summarized in Table 1. The second line of the table 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.

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
β = (1, -1) & α = (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.

Finally, the last line contains the results 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:

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

Since the long-term rate is weakly exogenous for all countries, 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 the 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. To conserve space - and since we are not really interested in the actual estimation results - Table 2 reports important diagnostic information only.

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 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). 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).

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 in the following as the rate of change of the conditional variance*.

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

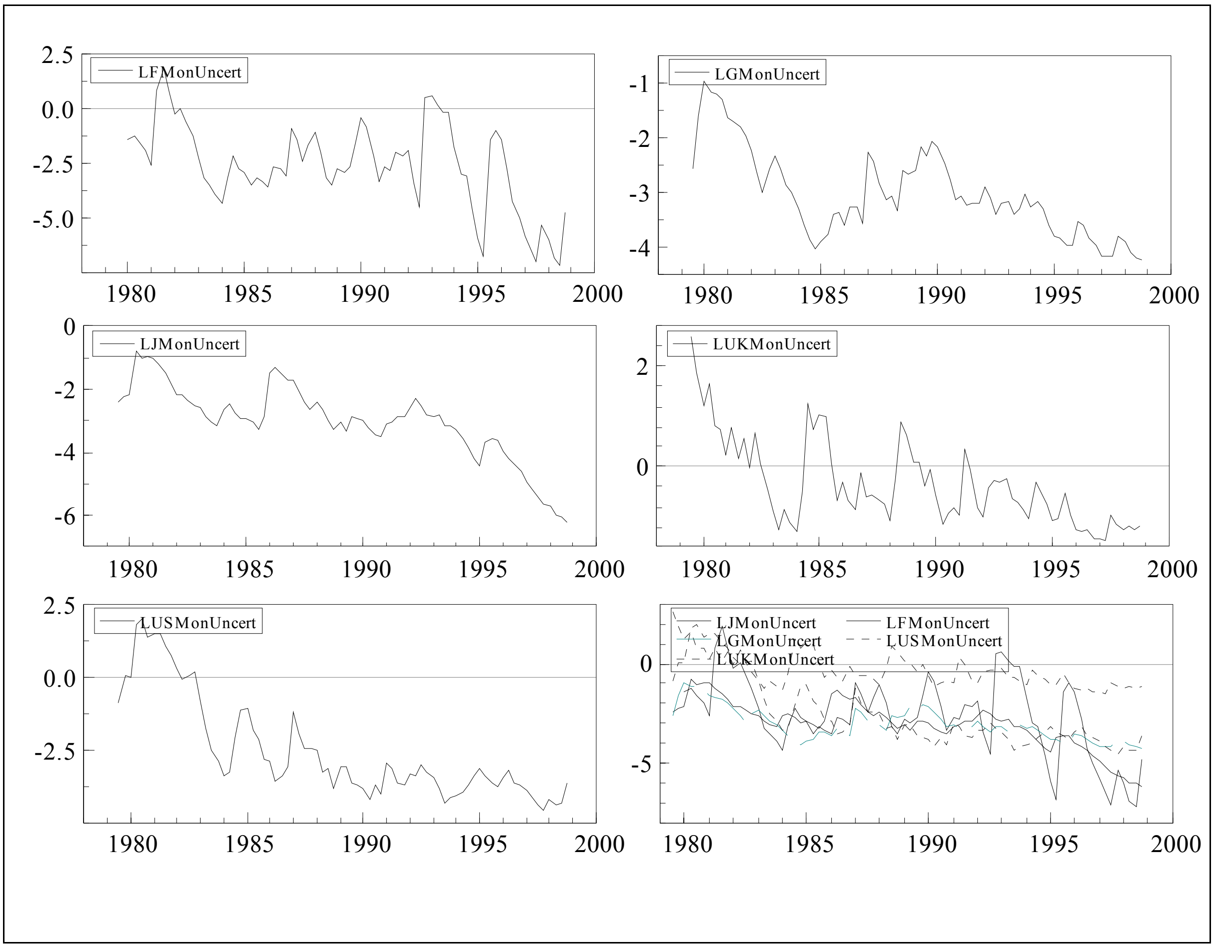

Figure 2 plots this measure of monetary uncertainty over time. It is apparent that the series are now stationary. Comparing the indicators for monetary uncertainty shows that the amplitude of the fluctuations generally declines over time with the exception of France. As the last graph shows, the series for France displays large fluctuations even at the end of the sample period.

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

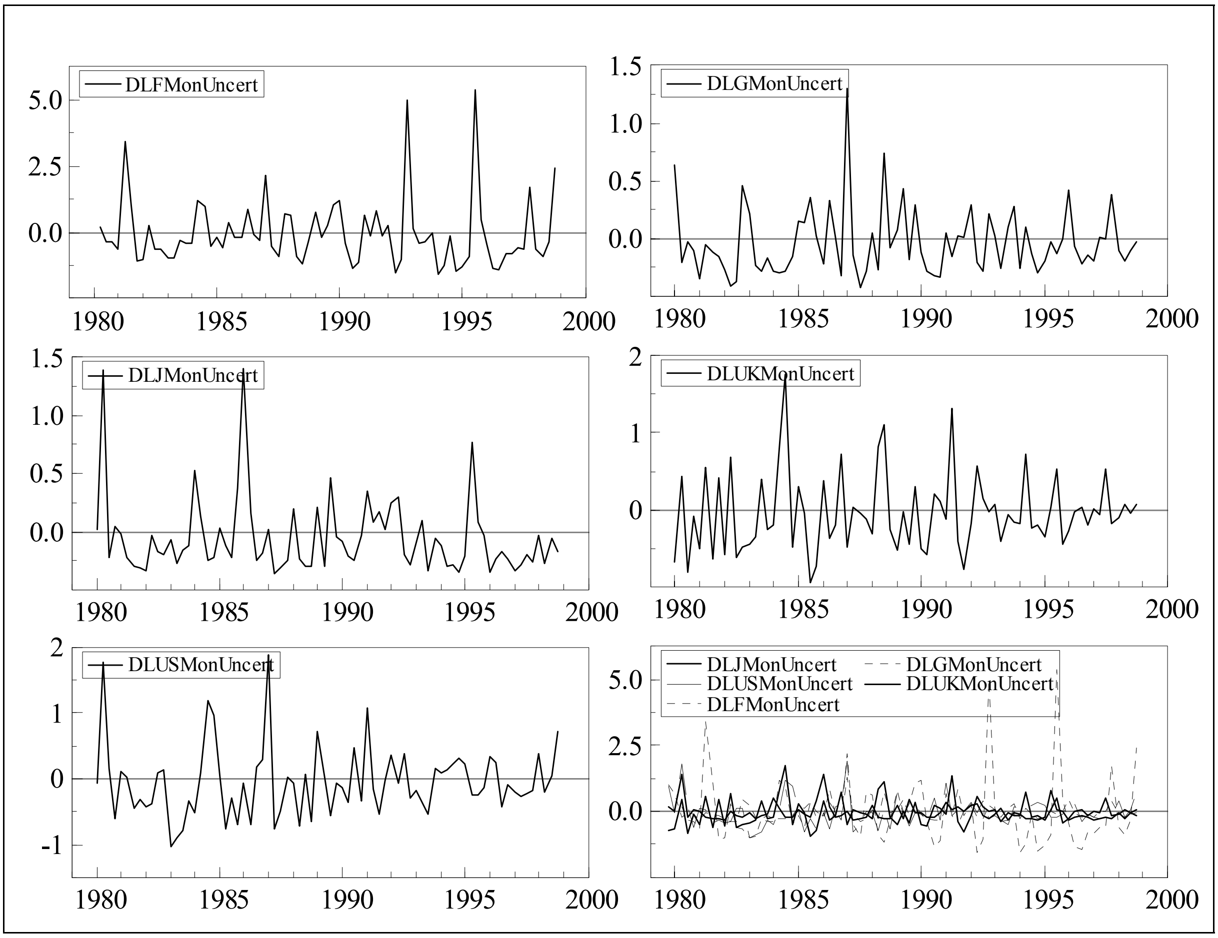

Table 3 provides some descriptive statistics of the indicator for monetary policy uncertainty.

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

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 largest. The UK and the US are intermediate cases, while Germany and Japan show the lowest variations. The highest correlation exists between France and Germany, which may be a reflection of 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 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.

It is unlikely that all of the conditional variance captured by our model can be directly attributed to monetary policy. We would maintain, though, that monetary policy can contribute to this uncertainty. If uncertainty per se affects wage setting, and monetary policy can influence uncertainty, then monetary policy will have an impact on wage 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. All series 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.

We model wage inflation using a Phillips-curve framework. Moreover, the actual specification proceeds within 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. To ensure that the estimated models are reasonably stable over time, 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.

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.

The only rejection of a null hypothesis in the out-of-sample Chow-test statistics occurs in the case of Japan. Here the main culprit appears to be the deflationary economic environment in Japan over the last years, which presumably causes the predictive failure.

Table 5 contains the estimated results of the simplified models. Most models display only few significant effects.

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

	France	Germany	Japan	UK	US
$\Delta Wage_{t-1}$		-0.257* (0.122)		0.497** (0.119)	0.572** (0.090)
$\Delta Wage_{t-2}$	0.089 (0.082)		0.447** (0.105)		
$\Delta Wage_{t-3}$	0.339** (0.080)				
$\Delta Wage_{t-4}$	0.290** (0.070)				
$NAIRU_t$			-0.004(*) (0.002)	-0.002(*) (0.001)	-0.0006 (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* (0.0004)				-0.0009 (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* (0.001)	0.013** (0.001)	0.004** (0.001)	0.008** (0.002)	0.006** (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. Due to the computation of out-of-sample tests, the actual estimation period ends in 1996:4. **, *, (*) indicate significance at a 1%, 5% and 10% level, respectively.

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.

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.[3] 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.

5. Robustness tests

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. Thus, the impact of uncertainty on wages is stable over time.

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 these variables as a robustness check. We test two indicators for future inflation: A contemporaneous *expected inflation rate* derived from an instrumental variable regression with four lags of inflation and the contemporaneous value plus four lags of

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

the output gap, and the contemporaneous *output gap* plus four lags, which can be interpreted as an indicator for marginal costs. To capture the impact of productivity, the current change in labor productivity and four lags are tested against the models in Table 5.

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. The output gap variables are not significant ($F(5,43) = 1.73$) and neither are the labor productivity indicators ($F(5,43) = 1.13$). In the German model, the instrumented current inflation rate is not significant. However, the output gap variables are significant at the 10% level ($F(5,56) = 2.07$). It turns out that the third lag of the German output gap is significantly negative at the 10% level. This result is neither consistent with the typical interpretation of the GDP gap as an indicator of marginal costs nor does it affect the estimate of our indicator of monetary policy uncertainty. The labor productivity variables are not statistically significant. ($F(5,56) = 1.1625$).

Turning to the case of Japan, we find again that the current instrumented inflation is not significant. The same conclusion holds when including the current output gap and four lags ($F(5,55) = 1.36$). While the four lags and current labor productivity are not significant as a group, including current labor productivity only leads to a significantly positive outcome ($F(1,60) = 4.85$). Including this variable in the model reduces the p-value for monetary uncertainty to 0.052 but cannot improve on the stability properties of the model as tested by the out-of-sample Chow-test. In the UK, the output gap variables are not significant ($F(5,48) = 1.05$) and neither is the current instrumented inflation rate ($F(1,52) = 1.95$). The labor productivity variables are not significant as a group. However, the current value is significant at the 10% level, and it enters the equation with a positive sign. None of these additions to the UK base model leads to a significant impact of monetary policy uncertainty on wage inflation.

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. Highly significant are the groups of output gap variables (F(5,55) = 6.65**) and productivity variables (F(5,55) = 2.98*). While the output gap variables, in particular current and first lag, show reverse signs, the current and lagged changes in labor productivity have a positive impact on wage inflation. However, none of these changes lead the coefficient on monetary uncertainty to become significantly negative.

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.

Fourthly, one may question the adequacy of the expectation theory of the term structure as a good model for explaining short-term interest rates. An alternative would be the use of a *Taylor rule as an explicit model of interest rate setting*. Clarida et al. (1998) explain current interest rates by a lagged value, the current output gap, and the inflation rate one year ahead. The estimates for the G5 countries derived by these authors are applied to our sample period employing monthly data on short-term interest rates, annualized inflation, and industrial production. We remove the part of the variation from our interest rate series that can be explained by their short term estimates for the Taylor rule (without the constant). Second, we fit an GARCH(1,1) model with a constant in the mean equation to the interest rate residuals. It turns out that in all cases we find significant estimates for one or both of the main GARCH(1,1) parameters (α_1 and β_1). Third, we store the conditional variance associated with the ARCH models and, as before, use the change in the logarithm of these conditional variances as the indicator for monetary policy uncertainty. Finally, we re-estimate the models in Table 5 substituting the indicator for monetary policy uncertainty indicator based on the interest rate expectation hypothesis by this alternative.

To conserve space, Table 6 shows only the estimates for the monetary policy uncertainty indicators within the final models (omitted results available upon request).

Table 6: Impact of monetary policy uncertainty derived from Taylor rules on wage inflation

	France	Germany	Japan	UK	US
Uncertainty_{t-i}	-0.002* (i =2)	-0.004* (i =4)	-0.002** (i =3)		
	(0.001)	(0.002)	(0.001)		

Notes: The model for Japan contains two impulse dummies (1981:1 and 1990:2) to remove non-normality from the residuals. As the model for France shows significant evidence of heteroscedasticity and autocorrelation, we have applied standard errors based on Andrews (1991).

We observe significantly negative effects of monetary policy uncertainty on wage inflation in France, Germany, and Japan, while we do not find significant effects in the case of the UK and the US. Thus, while there are slight changes in the dynamics of the models, we conclude that the core results outlined above are robust to the change in computing the uncertainty variable. In other words, our results are not particularly sensitive with regard to the specification of the base model for explaining short-term interest rates. However, we prefer the interest rate expectations hypothesis as a base model for constructing the monetary policy uncertainty indicator to the one using on Taylor rule estimates. Empirically, the cointegration relationship between long- and short-run interest rates appears to be much more robust than Taylor rules estimated by instrumental variable techniques. To summarize, the results of our models hold up quite well against these four types of robustness tests.

6. Conclusion

Extending an approach by Sorenson (1991), we develop a model that links wage setting behavior and monetary policymaking in unionized economies. Monetary policy uncertainty may have a disciplinary impact on wage growth in countries where labor unions internalize the influence of their actions on the monetary policy of the central bank. We test the model for five countries, Germany, France, Italy, the UK and the US by including an indicator of monetary policy uncertainty into a dynamic wage equation. Our empirical analysis is consistent with the view that monetary policy uncertainty may lead to wage restraint conditional on the existence of a relatively centralized behavior of the trade unions.

This finding has important consequences for the design of monetary policy institutions as well as 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 lower 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.

Further empirical research could further 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 encompassing 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.

Price data (monthly series)

CPI in IMF International Financial Statistics

Industrial production (monthly series)

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