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WAGE DYNAMICS NETWORK

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Wage Dynamics Network

This paper contains research conducted within the Wage Dynamics Network (WDN). The WDN is a research network consisting of economists from the European Central Bank (ECB) and the national central banks (NCBs) of the EU countries. The WDN aims at studying in depth the features and sources of wage and labour cost dynamics and their implications for monetary policy. The specific objectives of the network are: i) identifying the sources and features of wage and labour cost dynamics that are most relevant for monetary policy and ii) clarifying the relationship between wages, labour costs and prices both at the firm and macro-economic level.

The WDN is chaired by Frank Smets (ECB). Giuseppe Bertola (Università di Torino) and Julian Messina (Universitat de Girona) act as external consultants and Ana Lamo (ECB) as Secretary.

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The paper is released in order to make the results of WDN research generally available, in preliminary form, to encourage comments and suggestions prior to final publication. The views expressed in the paper are the author's own and do not necessarily reflect those of the ESCB.

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CONTENTS

Abstract

In this paper we present an extension of the Taylor model with staggered wages in which wage-setting is also influenced by reference norms (i.e. by benchmark wages). We show that reference norms can considerably increase the persistence of inflation and the extent of real wage rigidity but that these effects depend on the definition of reference norms (e.g. how backward-looking they are) and on whether the importance of norms differs between sectors. Using data on collectively bargained wages in Austria from 1980 to 2006 we show that wage-setting is strongly influenced by reference norms, that the wages of other sectors seem to matter more than own past wages and that there is a clear indication for the existence of wage leadership (i.e. asymmetries in reference norms).

Keywords: Inflation Persistence, Real Wage Rigidity, Staggered Contracts, Wage Leadership

JEL Classification: E31, E32, E24, J51

Non-Technical Summary

Standard models with nominal wage rigidities typically imply less inflation persistence than can be found in empirical data (cf. Roberts, 1995; Fuhrer and Moore, 1995; Chari et al., 2000). A common way to deal with this problem is to assume that a fraction of firms use rule-of-thumb behavior or automatic indexation of their prices (cf. Christiano et al., 2005; Smets and Wouters, 2007). These assumptions are, however, rather controversial since the existing microeconomic evidence on price-setting is not in line with this kind of behavior.

In this paper we analyze whether the process of wage-setting might be responsible for the observed excessive persistence and the strong influence of backward-looking behavior. The argument is based on the well-documented evidence from survey studies (cf. Bewley, 1999; Agell and Bennmarker, 2007) that wage-setting is considerably influenced by benchmark values that are typically related to past wage levels (e.g. to the workers' own past wages or to wages that have been previously set in other sectors of the economy). This assumption of benchmark values (or reference norms) thus provides a plausible and empirically validated way to make nominal variables more persistent.

The paper investigates this argument from both a theoretical and an empirical angle. In the theoretical part we extend the classical, two period staggered-wage model by Taylor (1980) to allow for reference norms. We show that the introduction of reference norms can considerably increase inflation persistence in this standard framework. In a further step we also demonstrate that the size of this impact depends on the precise type of the prevalent reference norm and on whether the importance of reference norms differs between the sectors of the economy. We discuss several different reference norms that correspond to frequently made assumptions in the literature. As our benchmark specification we use the assumption that the wage-setters are influenced by an "external norm", i.e. they look at the last wage level that has been set in the other sector. Different assumptions about reference norms, however, have an effect on the exact amount of additional persistence. A particularly important case in this context is the assumption of wage leadership where the wage in one sector works as a reference norm for the other sector but not vice versa. A structure of wage leadership is often said to be a good description of industrial relations in a number of Scandinavian and continental European countries. We show that such asymmetries in the importance of reference norms between sectors reduce the extent of persistence considerably.

In the empirical part of the paper we use a unique dataset on collectively bargained

wages in Austria from 1980 to 2006 to study the interplay between reference norms and staggered wages. The dataset comprises around 100 individual wage-setting units covering almost the complete national labor force. The data are well-suited to study these relations since the process of Austrian wage-setting shows a clear structure of staggering and also because it is often argued that the metal sector acts as a wage leader for the sectors that follow. The data enable us to construct the reference norms that were discussed in the theoretical part of the paper. The external reference norm is thus defined as the average wage increase of all other units since the last wage settlement of the own unit whereas the leadership norm is simply the wage increase in the metal sector. Moreover, we also construct a "habit norm" (defined as the last increase of the own wage rate) and a "standard of living" reference norm. Our main other explanatory variables for wages are forecasts for real economic activity and inflation.

Our results indicate that reference norms are in fact an important factor for Austrian wage settlements. By and large, the coefficient of the reference norm is slightly higher than the coefficient of forecasted inflation. Of all reference norms considered, the wage leadership norm performs best in the empirical analysis. This is based on parameter restrictions derived from the theoretical model, on model specification tests as well as on the results from a random coefficient model. In particular, we find that the wage-leading sector reacts significantly stronger to the general macroeconomic outlook than the following sectors.

Our findings have a number of consequences concerning inflation persistence and real wage rigidity. First, reference norms are an important factor in wage-setting and they need to be included in realistic and comprehensive dynamic macroeconomic models. Second, in order to get a complete picture about the sources and consequences of nominal rigidities it is important to study the microstructure of wage setting. Economies that are characterized by differences in the importance of reference norms between sectors will show less persistence than countries with a symmetric structure. Third, underlying differences in reference norms could also be responsible for the observed cross-country differences in inflation persistence and wage rigidities.

1 Introduction

Numerous studies have documented that standard models with wage rigidities (like the models by Taylor [1980] and Calvo [1983]) do produce considerably less endogenous persistence in output and inflation than can be found in empirical data (cf. Roberts, 1995; Fuhrer and Moore, 1995; Chari et al., 2000). The most common way to deal with this problem is to assume that there exists some amount of additional intrinsic persistence in prices. This is motivated, e.g., by the existence of "rule-of-thumb" price setters (cf. Galí and Gertler, 1999) or by the assumption that all firms that do not reoptimize simply index their prices to past inflation (cf. Christiano et al., 2005; Smets and Wouters, 2007). These are useful short-cuts that help to bring the standard models with nominal rigidities closer in line with the main properties of the observed data. Nevertheless, the assumptions of automatic indexation and rule-of-thumb behavior are also controversial and they have been criticized as being ad-hoc, implausible and at variance with the available evidence on actual price-setting behavior of firms (cf. Mankiw and Reis, 2002; Rudd and Whelan, 2006).[1]

Given this controversial role of backward-indexation in price-setting it seems natural to ask whether instead the process of wage-setting could be responsible for the observed excessive persistence. In fact, there exists an extensive survey literature documenting that wage-setting is affected by many more factors than can usually be found in standard labor market models (cf. Campbell and Kamlani, 1997; Bewley, 1999; Agell and Bennmarker, 2007). In particular, it has been shown that actual wage-setting policies are considerably influenced by benchmark values, e.g. by the workers' own past wages or by the wages in other sectors of the economy. Paying workers less than their benchmark value might cause motivational problems with detrimental effects on morale and productivity. Interestingly, these arguments from the labor market literature are only rarely incorporated into the commonly used dynamic macroeconomic models. There exist some papers that take up this line of thought but they differ in their focus and point of departure and they have also not been clearly connected to the issue of backward indexation. Among these papers some use models with relative wages (Buiter and Jewitt, 1981; Fuhrer and Moore, 1995; Ascari and Garcia, 2004) while others are based on efficiency wages (Danthine and Kurmann, 2004) or wage norms (Hall, 2005; Gertler and Trigari, 2006). The common thread in this

[1] E.g. "Backward indexation of prices, an assumption which, as far as I know, is simply factually wrong, has been introduced to explain the dynamics of inflation. And, because, once they are introduced, these assumptions can then be blamed on others, they have often become standard, passed on from model to model with little discussion" (Blanchard, 2008, 25).

literature is that wage-setters are assumed to have (explicit or implicit) benchmark wages or — as we will call them henceforth — reference norms that introduce an element of backward-looking behavior and thereby contribute to real wage rigidities.

The paper contributes to the literature from both a theoretical and an empirical angle. In the theoretical part we extend the classical, two period staggered-wage model by Taylor (1980) to allow for reference norms. In the benchmark case the reference norm is specified as an "external norm", i.e. wage-setters are assumed to look at the last wage level that has been set in the other sector. We show that the introduction of this reference norm can considerably increase inflation persistence. In a further step we also demonstrate that the size of this impact depends on the precise type of the prevalent reference norm and on whether the importance of reference norms differs between the sectors of the economy. It is crucial to take these issues into account since the empirical literature suggests that there exist sizable differences along these lines both within and between countries. Campbell and Kamlani (1997) and Bewley (1999), e.g., report that US workers mainly compare their wage rate with their own past wages and with the wages of other workers within the same firm ("internal reference norms"). Agell and Lundborg (2003) and Agell and Bennmarker (2007), on the other hand, document a much larger role for external norms in Sweden and considerable differences between Swedish and US survey data. In our dynamic model the amount of persistence turns out to depend on the assumption about the structure of reference norms, in particular on their degree of "backward-lookingness" and on the importance of external (i.e. cross-sectional) comparisons.

As far as the possibility of asymmetric reference norms between sectors is concerned it is important to note that it is frequently argued that a structure of "wage leadership" is present in a number of Scandinavian and continental European countries.[2] In such a system a specific sector (mostly the metal sector) acts as a wage leader that sets its wage levels more or less irrespective of what happens in the rest of the economy while the other sectors will take this wage rate as their reference norm (cf. Smith, 1996; Lindquist and Vilhelmsson, 2006; Traxler et al., 2008). Despite the empirical importance we know of no paper that has studied wage leadership in the framework of standard dynamic macroeconomic models. The set-up of our model allows us to tackle this issue. We show that asymmetries in the importance of reference norms between sector reduce the extent of persistence. This means that for two economies that are characterized by an identical structure and an identical average importance of reference norms the economy

[2] In the literature one can find a number of synonymous expressions for this phenomenon like "pay leadership", "wage spillovers", "pattern bargaining" or "key bargaining."

with asymmetric norms (e.g. wage leadership) will exhibit less persistence.

In the second part of the paper we then proceed to analyze whether the implications of the model are more than just a theoretical possibility. An empirical test of the interplay between reference norms and staggered wages requires a special kind of data that are not easily available for many countries. In particular, one needs a comprehensive set of wage data that clearly indicate the point of time when each wage rate has been set and for how long it has been valid. For this purpose we have been able to construct a unique dataset on collectively bargained wages in Austria from 1980 to 2006 that comprises around 100 individual wage-setting units covering almost the complete national labor force. These data are particularily apt for our question at hand since the process of Austrian wage-setting shows a clear structure of staggering. Employers and employees in the metal sector traditionally start their negotiations after the summer ("autumn bargaining round") while the wage-setters in other sectors follow until May of the following year. In order to study the importance of reference norms the use of these data is even more interesting since it is often argued that in Austria the metal sector acts as a wage leader for the following negotiations. We can use our dataset to actually test for this "folk wisdom" and to contrast it with other assumptions about the structure of reference norms.

To this end we construct different reference norms that follow the suggestions in the literature. Our results indicate that reference norms are in fact an important factor for Austrian wage settlements. By and large the coefficient of the reference norm is slightly higher than the coefficient of forecasted inflation. Furthermore, our findings clearly suggest that the wage rates of other units (the wage rate of the metal sector and/or of all other wage settlements since the last negotiations) matter more than internal habits (i.e. the own last settlement). Finally, we get strong indication of asymmetries and wage leadership in Austrian wage-setting behavior. This conclusion is supported by direct nested and non-nested statistical tests as well as by tests that are based on implications of the theoretical model. It is also suggested by an analysis of the individual, wage-unit specific coefficients that result from the estimation of a random coefficient model. In particular, we find that the wage-leading sector reacts significantly stronger to the general macroeconomic situation than the following sectors. The latter take the wage agreement of the metal sector as a guideline and thus put less emphasis on macroeconomic data and forecasts.

Our findings have a number of consequences concerning inflation persistence and real wage rigidity. First, reference norms are an important factor in wage-setting and they should be included in realistic and comprehensive dynamic macroeconomic models. The assumption of reference norms leads to similar reduced forms as the approaches that assume backward-looking price-setters, while at the same time being more in line with empirical evidence. Second, in order to get a complete picture about the sources and consequences of nominal rigidities it is important to study the microstructure of wage-setting. Economies that are characterized by differences in the importance of reference norms between sectors will show less persistence than countries with a symmetric structure.[3] Third, underlying differences in reference norms could also be responsible for the observed cross-country differences in inflation persistence and wage rigidities (cf. Cecchetti and Debelle, 2006; Dickens et al., 2007).

The paper is organized as follows. In the next section we present a simple model with staggered wages that allows for reference norms. In section 3 we describe our data on collective wage-bargaining in Austria and we report the results of the empirical analyses. Section 4 concludes.

2 A model with wage staggering and reference norms

2.1 The set-up

We use a variant of the Taylor (1980) model with staggered wage contracts that considers the role of reference norms in wage-setting. The model assumes that the total workforce is divided into two sectors of equal size where sector A negotiates the wage in periods $t = 0, 2, 4, ...$ while sector B negotiates in periods $t = 1, 3, 5, ...$ All wage contracts are fixed for two periods (where one period corresponds to half a year). The fixed length of wage contracts captures a prevalent feature of wage-setting in many countries (see section 3.2 below) and seems more appropriate than the assumption of Calvo contracts. The **wage-setting equation** (for the adjusting sector i) is assumed to take the following form:

$$w_t^i = \left(1 - \mu^i\right) \tilde{w}_t^i + \mu^i rn_t^i \tag{1}$$

where $\tilde{w}_t^i$ is the "pure wage target" of wage-setters in sector i , rn_t^i stands for their reference norm and μ^i is the relative weight of these two magnitudes. The pure wage target follows the specification in Taylor (1980), i.e.:

$$\tilde{w}_t^i = bp_t + (1-b)E_t p_{t+1} + \gamma^i \left(by_t + (1-b)E_t y_{t+1}\right) \tag{2}$$

[3] Carvalho (2005) and Dixon and Kara (2007) study the consequences of asymmetries in the contract length on the amount of persistence in Taylor models.

The pure wage target $\tilde{w}_t^i$ depends on the expected price level and the expected level of real activity (or excess demand) during the duration of the contract. The expected price level is given by $(bp_t + (1-b)E_t p_{t+1})$, where p_t is the price level in period t and b is the relative weight of the two periods over which the contract is valid. Similarly, y_t is the measure of aggregate demand in period t and γ^i represents the degree of "real rigidity" (cf. Ball and Romer, 1990). The higher is γ^i, the stronger the wage target $\tilde{w}_t^i$ reacts to the conditions in the real economy and thus the smaller is the degree of real rigidity.[4]

If $\mu^i = 0$ wage-setters do not have reference norms and the wage-setting equation (1) coincides with the original formulation in Taylor (1980). For $\mu^i > 0$, however, wage-setting is assumed to be influenced by reference norms. This specification is based on the observation that in their negotiations wage-setters typically also care about other nominal variables in a direct fashion (and not only via their impact on current and future price levels). The reasons for such a behavior can be manifold and various explanations have been proposed in the literature that range from efficiency wage and relative wage models (Fuhrer and Moore, 1995; Danthine and Kurmann, 2004; Ascari and Garcia, 2004) to models with wage norms (Hall, 2005; Gertler and Trigari, 2006). Although these models differ in their motivation and also in their particular specification they typically imply an additional role for past wage rates. The wage-setting equation (1) is meant to capture the common element of these different specifications.

For our benchmark model we use the straightforward assumption of **"external norms"** (i.e. norms that refer to the wage in the other sector):

$$rn_t^i = w_t^{-i} = w_{t-1}^{-i}, \tag{3}$$

where $(-i)$ stands for the other sector. Equation (3) thus says that wage-setters in sector A look at the wage level in sector B $(rn_t^A = w_t^B = w_{t-1}^B)$ while wage-setters in sector B look at the wage rate in sector A $(rn_{t+1}^B = w_{t+1}^A = w_t^A)$. This is a reasonable specification that is in line with the results from survey studies (e.g. Agell and Bennmarker, 2007). Other reference norms will be discussed below. Note that for the extreme case with $\mu^i = 1$ assumption (3) implies complete stickiness of wages (i.e. $w_t^i = w_{t-1}^i = w_{t-2}^i$).

[4]Taylor did not derive his wage-setting equation from "first principles" but he motivated it as being "simple and plausible" (Taylor, 1980, p. 4). It can be shown, however, that a system of equations that is very similar to the ad-hoc specification of the Taylor model can be derived as the linearized solution to a fully fledged intertemporal optimization model (see appendix A and Ascari, 2000; Huang and Liu, 2002). In this case all the variables have to be interpreted as being percentage deviations around their respective steady states.

Equations (1) to (3) describe the basic structure of wage-setting in our framework that will underlie our empirical estimations. In the following we want to use this description of wage-setting behavior in an otherwise standard dynamic model in order to show how the existence of reference norms changes the persistence of inflation. To this end we have to specify how the other endogenous variables are determined.

Prices are set as a **mark-up over wages** and the price level p_t is thus given by the average wage:

$$p_t = \frac{1}{2}\left(w_t^A + w_t^B\right) \tag{4}$$

As in many versions of the Taylor model we also assume that **aggregate demand** y_t depends on nominal demand (or the money supply) m_t and the price level p_t.[5] In particular:

$$y_t = m_t - p_t \tag{5}$$

In order to close the model we have to make an assumption about the determination of monetary policy. Since we are primarily interested in how the amount of *intrinsic* persistence changes across different wage-setting regimes this assumption is not crucial and we thus use the simple specification that m_t follows an autoregressive process:

$$m_t = \rho m_{t-1} + \nu_t, \tag{6}$$

where ν_t is an i.i.d. error term. This completes the description of the model.

2.2 The solution of the model

In appendix B we report the solution for the general case of the dynamic model with $\mu^A \neq \mu^B$, $\gamma^A \neq \gamma^B$ and $b \neq \frac{1}{2}$. We show that this solution can be written as:[6]

$$w_t^i = \lambda^i w_{t-1}^{-i} + \theta^i m_t \tag{7}$$

As is shown in appendix B the root λ^i that captures the amount of intrinsic persistence between periods can differ across the two sectors, i.e. $\lambda^A \neq \lambda^B$. One can insert, however,

[5]This is done in Taylor's original model (1980) and also in various later contributions to this topic (e.g., Chari et al., 2000; Karanassou and Snower, 2007). Other papers, especially in the context of the "New Keynesian Phillips Curve", treat y_t as an exogeneous forcing variable (e.g., Roberts, 1995; Galí and Gertler, 1999).

[6]Note that equation (7) is only valid for the periods where the wage in sector i is changed. In the other periods wages are fixed and it thus holds that $w_{t+1}^i = w_t^i$.

$w^{-i}_{t-1} = \lambda^{-i} w^i_{t-2} + \theta^{-i} m_{t-1}$ into (7) to express the wage in sector i in dependence of its *own* last wage w^i_{t-2}, i.e.:

$$w^i_t = \lambda^i \lambda^{-i} w^i_{t-2} + \theta^i m_t + \lambda^i \theta^{-i} m_{t-1} \tag{8}$$

The level of *annual* persistence (i.e. from $t-2$ to t etc.) is thus the same in both sectors. We denote this degree of annual persistence by $\Lambda \equiv \lambda^A \lambda^B$ and we will use this measure for the numerical examples in the following sections.

In a symmetric world the two sectors differ only with respect to the period of time in which their wages are negotiated. In particular, this case is characterized by $\mu^A = \mu^B = \mu$, $\gamma^A = \gamma^B = \gamma$ and $b = \frac{1}{2}$. It is evident (and shown in appendix B) that in this situation the autoregressive persistence measure λ^i is the same in both sectors, i.e. $\lambda^A = \lambda^B = \lambda$. It can be calculated as:

$$\lambda = \frac{1 + \gamma + \mu(1-\gamma) - 2\sqrt{\gamma(1-\mu^2) + \mu^2}}{(1-\gamma)(1-\mu)}, \tag{9}$$

which implies an annual persistence measure of $\Lambda = \lambda^2$. In the absence of reference norms ($\mu = 0$) this model corresponds exactly to the formulation of the Taylor model that can be typically found in the literature (e.g. Romer, 2006, chap 6; Ascari, 2003; Karanassou and Snower, 2007). In particular, for $\mu = 0$ (9) reduces to the well-known expression:

$$\lambda = \frac{1 - \sqrt{\gamma}}{1 + \sqrt{\gamma}} \tag{10}$$

For both (9) and (10) it holds that λ decreases in γ ($\frac{\partial \lambda}{\partial \gamma} < 0$). The more strongly wages react to excess demand (the lower real rigidities) the lower will be the degree of persistence.

We can use the expression in (10) to briefly discuss the inflation or output persistence puzzle as it arises in the context of the Taylor model (cf. Fuhrer and Moore, 1995; Chari et al., 2000). The basis of the problem is that in microfounded models γ is not a free variable but rather depends on a number of structural parameters. In particular, under certain assumption one can show (see appendix A) that $\gamma = \frac{-\eta_{CC} + \eta_{LL}}{1 + \theta \eta_{LL}}$, where θ is the elasticity of substitution between different varieties of goods and η_{cc} and η_{ll} are the inverses of the intertemporal elasticity of substitution of consumption and of labor supply, respectively. Standard numbers for these parameters are (cf. Ascari, 2000, Table 1 or Dixon and Kara, 2007): $\eta_{cc} = -1$, $\eta_{ll} = 3.5$, $\theta = 6$. These values imply a value for the real rigidity of $\gamma = 0.21$. Using different acceptable parameter values (e.g. a higher elasticity of marginal consumption η_{cc} or a lower elasticity of labor supply η_{ll}) would suggest even higher values of $\gamma = 0.3$ or above.[7] Using (10) this corresponds to a degree of period-on-period persistence of $\lambda = 0.382$ (for $\gamma = 0.2$) and $\lambda = 0.292$ (for $\gamma = 0.3$) and to a degree of annual persistence of $\Lambda = 0.146$ and $\Lambda = 0.085$, respectively. These degrees of persistence appear, however, too low and to be in contradiction to the existing empirical literature. In particular, empirical estimates typically show values for the degree of annual persistence that are larger than 0.25 and often larger than 0.5 (see e.g. Jeanne, 1998). In a similar fashion Ascari (2003) "defines a significant degree of persistence to be a value of λ of at least 0.5" (p. 526). We can use this benchmark value of $\lambda = 0.5$ or ($\Lambda = 0.25$) in order to judge whether a model is successful in producing an empirically plausible degree of persistence for values of γ that are in line with the standard structural parameters (i.e. between $\gamma = 0.2$ and $\gamma = 0.3$).

In fact, the expression for λ in (9) indicates that the introduction of reference norms increases the amount of intrinsic persistence considerably. This is illustrated in Figure 1 for the measure of annual persistence Λ. For both $\gamma = 0.2$ and $\gamma = 0.3$ an increase in the importance of reference norms to values between $\mu = 0.2$ and $\mu = 0.5$ is sufficient to increase Λ to levels that are in line with the empirically observed degree of persistence. Note also that $\lim\limits_{\mu \to 1} \lambda = 1$. This means that under the assumption of external reference norms (cf. (3)) we can have complete persistence ($\lambda = 1$) if comparisons play an overwhelming role in wage-setting.

`Insert Figure 1 about here`

2.3 The role of different reference norms

We have seen in the last subsection that the introduction of reference norms can increase inflation persistence considerably. In this and in the next subsection we want to study how this impact changes once we allow for different references norms and for possible asymmetries between sectors in the importance of reference norms.

In the benchmark specification (3) we have assumed that wage-setters have backward-looking external reference norms. In the related literature, one can however find a number of different assumptions concerning the variable(s) that primarily influence wage-setting

[7] The assumption of decreasing returns to scale in production ($\alpha < 1$) does not change the main message. In particular, in this case we get that (see Ascari, 2003): $\gamma = \frac{-\eta_{cc} + \frac{1}{\theta(1-\alpha)+\alpha}\eta_{ll}}{1+\frac{\theta}{\theta(1-\alpha)+\alpha}\eta_{ll}}$. Using $\alpha = \frac{2}{3}$ and again $\eta_{cc} = -1$, $\eta_{ll} = 3.5$ and $\theta = 6$ we get $\gamma = 0.26$. In a model with staggered price-setting Chari et al. (2000) get a value for γ that is even larger than 1 (since they have $\gamma = \eta_{ll} - \eta_{cc}$).

decisions (see Agell and Bennmarker, 2007; Danthine and Kurmann, 2006). The choice of the reference norm can evidently have an important impact on inflation persistence.

The most direct way to see this is if we assume that instead of (3) wage-setters only make **contemporaneous comparisons**, i.e.

$$rn_t^i = w_t^i \tag{11}$$

Inserting this assumption into (1) implies that $w_t^i = \tilde{w}_t^i$ and we are back to the model without reference norms. In general, the impact of the reference norm on the stickiness of wages will depend on the degree of "backward-lookingness" and on the extent to which they are directed to the other sector. We want to illustrate this by using two alternatives to the external norm that are discussed in the literature. For the first alternative specification we assume a "**price indexation norm**" where reference norms are given by the actual price (or wage) level. This corresponds to an assumption that is frequently made in the related DSGE literature (cf. Christiano et al., 2005; Smets and Wouters, 2007). In these papers it is assumed that wages that are not optimally chosen in a certain period are simply indexed to the past or current rate of inflation. Transformed from growth rates to levels and translated into the model with reference norms this can be expressed as:

$$rn_t^i = p_t \tag{12}$$

For the second alternative we assume that wage-setters have "habits", i.e. their reference norm consists of the last increase of their own wage. This "**habit-persistence norm**" is thus given by:

$$rn_t^i = w_{t-2}^i \tag{13}$$

In Table 1 we compare the degrees of inflation persistence for the three alternative reference norms and for different parameter values for γ and μ. The numbers in the table correspond to the measure of annual persistence Λ since this allows for better comparisons once asymmetries between sectors are taken into account. For the habit persistence norm the solution cannot be expressed in an AR(1) form as in (7) and in this case the persistence measure corresponds to the sum of the first two lags.[8]

[8] One can think of many other possible reference norms. E.g., one could specify rn_t^i in such a way that it corresponds to the model by Fuhrer and Moore (1995) or to the alternative version of their model in Holden and Driscoll (2003). Also one could use a "forward-looking external norm" where wage setters take into account that the norm will be present today *and* in the next period (i.e. $rn_t^i = bw_t^{-i} + (1-b)E_t w_{t+1}^{-i}$). We stick here to the alternative norms (12) and (13) since they are most frequently discussed in the

Insert Table 1 about here

We observe that inflation persistence is lower for the price indexation norm and higher for the habit persistence norm. We have, e.g., that for $\gamma = 0.2$ and $\mu = 0.5$ the parameter Λ is 0.702 for the case of the external norm while it is only 0.539 for the price indexation norm but 0.916 for the habit persistence norm. For the price indexation norm this follows from the fact that the norm p_t now includes not only the past wages of the other sector w_{t-1}^{-i} but also wages w_t^i that are set in the own sector in the current period and thus include already a reaction to the current economic situation. Persistence is higher for the habit persistence norm than for the external norm since in this case there is no cross-sectional interaction in reference norms and this increases the amount of sluggishness. These results imply that, not surprisingly, the type of reference norm can have a significant impact on the degree of persistence. This is potentially important since differences in reference norms seem to be a real-world phenomenon. In fact, Agell and Bennmarker (2007) document considerable differences in reference norms between Sweden and the US where external norms are found to play a larger role for Swedish firms.[9]

2.4 Asymmetries in the importance of reference norms

The case of asymmetric importance of reference norms is also highly relevant since it captures the argument that there exist differences in the behavior of wage-setters across sectors. A particularly important example for such asymmetries is the case of wage leadership. In the language of our model this would mean that the wage setters in both sectors have external reference norms while $\mu^B > \mu^A$ (if sector A is the wage leader) and possibly $\mu^A = 0$. In Table 2 we report the values of Λ for two levels of γ ($\gamma = 0.2$ and $\gamma = 0.3$) and for different assumptions about the importance and possible asymmetries in external reference norms. In particular, we denote by $\bar{\mu}$ the average importance of reference norms in the economy, i.e. $\bar{\mu} = \frac{1}{2}\left(\mu^A + \mu^B\right)$. In the first column we show the results for $\bar{\mu} = 0$ and in the second block of columns the results for $\bar{\mu} = 0.5$ (both for a symmetric case with $\mu^A = \mu^B = 0.5$ and an asymmetric one with $\mu^A = 0$ and $\mu^B = 1$).

Insert Table 2 about here

literature.

[9] They explain this finding in the following way: "The precision of the information about external pay appears to be higher among workers in unionized firms" (Agell and Bennmarker, 2007, p. 363).

Table 2 shows that an increase in the importance of reference norms from $\bar{\mu} = 0$ to $\bar{\mu} = 0.5$ increases Λ from 0.146 to 0.702 (for $\gamma = 0.2$). If there are, however, asymmetries in reference norms ($\mu^A = 0$ and $\mu^B = 1$) then Λ is significantly lower at 0.5. This drop in Λ is even larger for $\gamma = 0.3$ where the increase in Λ from 0.085 (for $\bar{\mu} = 0$) to 0.602 (for $\bar{\mu} = 0.5$) is almost halved (to $\Lambda = 0.368$) for the case of asymmetric reference norms.

The results of Table 2 indicate that it is important to know if wage-setting is characterized by asymmetries, e.g. by an outright system of wage leadership. In the latter case, corporatist countries with a clear pattern of staggered wage-setting might still face a rather low level of persistence (even if the *average* importance of reference norms is large). This phenomenon could thus be partly responsible for the fact that different countries with apparently quite similar labor market institutions show considerably different degrees of inflation persistence (cf. Cecchetti and Debelle, 2006).

So far we have assumed that the degree of real rigidity γ^i is identical in the two sectors. If this is not the case then it can be shown that a sufficient condition for a system with wage leadership to have less persistence than a symmetric system is that the degree of real rigidity in the leading sector is smaller than the one in the following sector (i.e. $\gamma^A > \gamma^B$). This conforms to the often heard argument that a stable system of wage leadership presupposes that the wage leader is situated in a competitive segment of the economy and not, e.g., in the public sector.

An additional source of asymmetry is the potential clustering of wage contracts. In the empirical dataset that we are going to use later we see, e.g., that almost 50% of all new wage agreements take effect in the months January to March. This clustering can have important implications for the transmission of monetary policy (cf. Olivei and Tenreyro, 2007). For the question at hand, however, the effects are rather muted. Under the assumption that sector A subsumes only 10% of all firms (and $\gamma = 0.3$) the persistence measure Λ can be calculated as: 0.034 (for $\bar{\mu} = 0$), 0.601 (for $\bar{\mu} = 0.5$) and 0.372 (for $\bar{\mu} = 0$ and $\mu_A = 0$). These figures are close to the results for the case with symmetric sector sizes.

On the whole we can conclude that both the specific nature of the reference norms and asymmetries in the importance of the norms matter for the persistence of inflation.

3 Empirical Part

In order to empirically test the interplay between staggered wages and reference norms one needs a particular set of wage data that include information about the actual time of wage changes. We have been able to construct such a dataset that is based on collectively bargained wages in Austria. We use this dataset in the following to investigate the role of reference norms in an economy with a clear structure of wage staggering. We want to answer three questions: First, do wage-setters have reference norms or is their behavior only influenced by expectations about prices and aggregate demand as assumed in the standard model? Second, if reference norms do play a role, can we determine *which* formulation of reference norms is most appropriate? Third, is their any indication of asymmetries in wage-setting behavior ("wage leadership")?

3.1 Estimation Equation

The main equation for estimation follows directly from equation (1). If we take the first difference of equation (1) together with (2) we get (for $b = \frac{1}{2}$):

$$\begin{aligned} \Delta w_t^i &= (1-\mu^i)\frac{1}{2}\left\{\Delta p_t + E_t \Delta p_{t+1} + \gamma^i \left(\Delta y_t + E_t \Delta y_{t+1}\right)\right\} + \mu^i \Delta r n_t^i + \\ &\quad (1-\mu^i)\frac{1}{2}\left\{(p_t - E_{t-1} p_t) + \gamma^i \left(y_t - E_{t-1} y_t\right)\right\} \end{aligned} \tag{14}$$

Equation (14) states that wage growth in sector i will depend on expected inflation and expected changes in real activity over the duration of the contract. The expressions in the second line of (14) are expectational errors that should be zero on average if people form rational expectations (see Roberts, 1995). We can generalize this equation to a model with monthly staggering. In order to distinguish clearly between a year τ and a month j we denote with $w_{j,\tau}^i$ the wage that is set by the wage-setting unit i in month j in year τ. We denote by $\Delta\tilde{p}_{j,\tau}$ the rate of inflation over the upcoming year starting in month j. This means, e.g., that $\Delta\tilde{p}_{1,\tau} = \frac{1}{12}\left(\Delta p_{1,\tau} + \Delta p_{2,\tau} + ... + \Delta p_{12,\tau}\right)$, $\Delta\tilde{p}_{2,\tau} = \frac{1}{12}\left(\Delta p_{2,\tau} + ... + \Delta p_{12,\tau} + \Delta p_{1,\tau+1}\right)$ etc., where $\Delta p_{j,\tau} = (p_{j,\tau} - p_{j,\tau-1})$. In a similar fashion we denote the change in real activity over the next year by $\Delta\tilde{y}_{j,\tau}$. Our estimation equation takes the following form:

$$\Delta w_{j,\tau}^i = \beta_0^i + \beta_1^i E_{j,\tau}\Delta\tilde{p}_{j,\tau} + \beta_2^i E_{j,\tau}\Delta\tilde{y}_{j,\tau} + \beta_3^i \Delta r n_{j,\tau}^i + \left(\delta^i\right)' \mathbf{X}_{j,\tau}^i + \varepsilon_{j,\tau}^i, \tag{15}$$

where β_0^i is an individual effect and $\mathbf{X}_{j,\tau}^i$ is a vector of additional (possibly wage-unit-specific) regressors that might have an impact on individual wage growth and that are used in some of the following specifications. These additional regressors include, e.g., expectational error terms, time dummies and measures for the length of the wage contract. We have also introduced an error term $\varepsilon_{j,\tau}^i$ which may include sectoral or aggregate shocks. As written in equation (15), the empirical specification allows for possible heterogeneous reactions across wage-setting units. This refers not only to the constant β_0^i as in normal fixed or random effects models but also to the slope coefficients β_1^i to β_3^i and $(\delta^i)'$.

Furthermore, it is important to note that the theoretical model makes strong predictions about the size of the coefficients. Most importantly, the model implies that $\beta_1^i = (1-\mu^i)$ and $\beta_3^i = \mu^i$ and that a correct empirical specification should thus be consistent with the condition that $\hat{\beta}_1^i + \hat{\beta}_3^i = 1$. The close connection between the theoretical model and the estimation equation is thus quite helpful in analyzing and interpreting the results. In order to directly test the wage-setting equation (15) we have to find data for the main variables $\Delta w_{j,\tau}^i$, $\Delta\tilde{p}_{j,\tau}$, $\Delta\tilde{y}_{j,\tau}$, and $\Delta rn_{j,\tau}^i$. We will describe our data sources after giving a short overview of some specificities of the Austrian system of wage bargaining.

3.2 The Austrian system of collective wage bargaining

Wage determination in Austria is strongly dominated by a system of collective bargaining.[10] On the side of the employees there is a peak organization, the Austrian Federation of Trade Unions, to which the individual trade unions are attached. These individual trade unions are mainly organized along sectoral and occupational dimensions. On the side of the employers there is also a central organization, the Austrian Federal Economic Chamber, which covers basically all private companies. Collective bargaining mainly takes place at the sectoral and industry level and regional differences in wage agreements do not play an important role. Although collective bargaining is in principle confined to the private sector there is also an influential public sector union that negotiates over wages with representatives of the government. The coverage rate of the collective agreements is very high (around 95%). At the moment around 400 collective agreements are signed each year of which around 250 are national agreements. Many of these agreements, however, refer only to a small number of employees while around 20 large agreements together cover more than half of the complete labor force. Since the beginning of the 1980s the annual bargaining process follows a similar pattern where the metalworking industry starts the

[10] Details can be found in Traxler et al. (2001).

round of negotiations in September or October ("autumn bargaining round"). Since 1984 the collective agreement in this sector (which covers about 11% of the total wage bill) always takes effect in November. The other wage-setting units follow the metal sector in a staggered fashion. The collective agreement of the wholesale and retail trade sector that also represents a large part of the labor force (around 13%) normally becomes effective in January as is the case for the public sector. Finally, a number of important sectors typically have their wages only set in May (this is true, e.g., for construction and parts of the chemical and tourism sector which together cover around 10% of the wage bill).

3.3 The data

3.3.1 Collective wage agreements

Unfortunately, there does not exist an accessible complete database for all collective wage agreements in Austria. There exists, however, an "index of agreed minimum wages" that subsumes data for a large number of disaggregated wage-setting units that are mostly organized along sectoral lines. We use these disaggregated series to construct an annual dataset that contains for each included wage-setting unit i the annual increase in the collectively bargained wage $\Delta w^i_{j,\tau}$, the month in which this agreement came into effect and the duration until the next agreement is reached. We have to exclude some units either because we do not have data over the whole time span or because they refer to quite heterogeneous sectors with rather erratic patterns (e.g. many small changes each year). This leaves us with a number of 100 individual times series for collectively bargained wages comprising 92% of the total database. We focus on the time period from 1980 to 2006. Table 3 summarizes these data.[11] One sees that most contracts are signed in winter or spring (January to June). In fall only around 10% of new agreements are concluded but these include the important agreement of the potentially wage-leading metal sector. Most wage agreements ($> 80\%$) are valid for exactly one year.

`Insert Table 3 about here`

Our data on disaggregated wages have two potential drawbacks. First, the available data only indicate when a new collective agreement became effective and not when the change was negotiated. It could in principle be the case that there is a longer time lag between a wage increase and the time when it has been scheduled. For the estimation of

[11] In appendix C we provide more details on the construction of our dataset.

equation (15) it is important to know on which macroeconomic forecasts the wage-setters could have based their decisions. Casual observations as well as personal information by involved experts suggest, however, that the time lags between the end of the negotiations and the implementation are rather short. Second, the available time series only report the collectively bargained increase in the minimum wage for each unit. As in many other countries, in Austria effective wages are often higher than these agreement-specific minimum wages. Although this is admittedly a handicap of the dataset it is probably less severe than one could expect. The increase in *effective* wages which are also negotiated in some industries is mostly parallel to the increase in minimum wages. Furthermore, the development of the collective wage index follows closely the one of the comprehensive variable "compensation per employee for the total economy" from national accounts.

3.3.2 Macroeconomic forecasts and forecast errors

For the expected values $E_{j,\tau}\Delta\tilde{p}_{j,\tau}$ and $E_{j,\tau}\Delta\tilde{y}_{j,\tau}$ in (15) we use the quarterly forecasts of the Austrian Institute of Economic Research (WIFO). This institute has a long tradition in forecasting the future path of the Austrian economy and its results are widely published in the media and are also the "official" numbers that are used in the collective wage negotiations. The forecasts are typically published in March, June, September and December each year and they include forecasts for the current and the next year for a number of macroeconomic variables. We use the figures for the growth rate of GDP, for the rate of inflation and for the unemployment rate. In order to match these forecasts with the time series of collectively bargained wages we assume that wages that come into effect in a certain month are based on the most recent forecasts available in the previous month.[12] The expected development of a variable over the duration of the wage contract is calculated as the weighted average of the forecasts for the current and for the next year.[13]

As far as the variable for real activity $E_{j,\tau}\Delta\tilde{y}_{j,\tau}$ is concerned there exists a long discussion on which is the most appropriate way to measure it (cf. Roberts, 1995; Galí and Gertler, 1999; Rudd and Whelan, 2006). In the related literature one can find specifications that use — among others — the output gap, real marginal costs, the labor share

[12]So strictly speaking, the information set of expectations formed in month j contains only information that was available at the end of month $j-1$.

[13]So we assume, e.g., that the wage agreement that became effective in May 2002 is based on an expected rate of inflation that is calculated as 8/12 of the WIFO-March-2002 forecast for the current year plus 4/12 of WIFO-March-2002 forecast for the next year. For details see appendix C.

and the unemployment rate. Due to problems with data availability we base our estimations on forecasts of GDP growth and the change in the unemployment rate. The first is an appropriate measure for real activity as it is closely related to changes in the output gap. The use of the change in the unemployment rate as a measure for business cycle conditions can be motivated by Okun's law as noted by Roberts (1995).

In some of our estimations we also use forecast errors as suggested by the theory that leads to the empirical specification in (15). The forecast error of a variable is measured as the difference between the realized and the expected value. For some of our robustness tests we will also use different macroeconomic aggregates (e.g. lagged instead of forecasted values). In all cases where we do not have monthly variables we interpolate them in the same way as it is done in the construction of monthly series for the forecasts.

3.3.3 Reference norms

We use 4 reference norms that have been discussed in section 2:

- **External reference norms** ("external norms"): For each wage-setting unit the reference norm is given by the (weighted) average increase of wages in all other units that could be observed since the last time that its own wage level had been changed.
- **Wage leadership reference norms** ("leadership norms"): Each wage-setting unit takes the wage increase in the metal sector as its reference norm. The wage-setters in the metal sector do not have reference norms. There are 14 wage-setting units which typically contract in November. We have subsumed 4 of these units into the category "wage-leader" since all of these units are attached to the metal sector, negotiate together and have reached almost perfectly correlated wage agreements. The remaining 10 wage-setting units in November include, e.g., units in the chemical and the paper industry.
- **Habit reference norms** ("habit norms"): Each wage-setting unit regards its own last wage change as the reference norm.
- **Price indexation norms** ("indexation norms"): All wage-setting units take the average inflation rate over the last year as their reference norm. This could be a reasonable assumption if, e.g., unemployment benefits are indexed to changes in the price level.

3.4 Empirical Strategy

It is not straightforward to say which reference norm is the correct assumption. The results of the survey studies indicate that various reference norms might be important. Furthermore, some of the norms are clearly not independent of each other. For example, even if all wage-units have a leadership norm the external norm will also show a significant influence due to the staggered structure of wage agreements. Therefore we will not a priori limit ourselves to a single norm but rather try to find out empirically which norm provides the best description of the data.

The choice of the right empirical specification to test our main hypotheses involves a number of issues. The most important one is related to the question whether it is reasonable to assume that all wage-setting units react to the macroeconomic variables and the reference norms in the same fashion. In fact, our theoretical model suggests that there might be differences in the weight of the importance of reference norms μ^i and possibly also in the degree of real rigidity γ^i. These differences should be reflected in heterogeneous regression coefficients as written in equation (15).[14] In order to allow for this possible heterogeneity we will estimate our benchmark equation not only with the usual homogeneous coefficients techniques (i.e. with fixed effects [FE] and random effects [RE] models) but also with two varying-coefficients estimation methods that can be frequently found in the literature: the random coefficients (RC) model by Swamy (1970) and the mean group (MG) estimator suggested by Pesaran and Smith (1995). A discussion of these methods can be found in Hsiao (2003, chap. 6) and in Hsiao and Pesaran (2004). The basic difference between the RC and the MG estimator is related to their assumption about the nature of heterogeneity. The MG estimator is based on the assumption that deviations from the mean coefficient are deterministic while the RC estimator assumes that they are stochastic. Accordingly, the MG estimator is defined as the simple average of individual OLS estimates while the RC estimator uses a weighted average of these estimates where the optimal weights are inversely proportional to the covariance matrices. Hsiao and Pesaran (2004) show that the two estimators are equivalent if the number of time periods is sufficiently large. As we will see below, our data strongly reject the assumption of homogeneous coefficients and we will mostly work with the RC specification.

A second specification issue concerns the treatment of possible common trends in

[14] This is confirmed in a different context by the results in Imbs et al. (2007) who show that homogeneity in pricing behavior is strongly rejected for French data.

the data (like a general decrease in the rate of inflation). This could suggest the use of annual time dummies or of other variables that capture these time trends. It is, however, not clear whether this is appropriate for our research question since general time trends in unit-specific wage growth should already be captured by the movements in the nominal variables on the right hand side of the wage-setting equation. Also, in the case of specifications with heterogeneous slope coefficients the inclusion of annual time dummies would be impossible. In order to allow for some broad changes in wage-setting over time the specifications include decade time dummies.[15]

Our analysis involves three steps. First, we will compare the average estimated coefficients and we will investigate whether and for which reference norm the theory-based condition $\hat{\bar{\beta}}_1 + \hat{\bar{\beta}}_3 = 1$ is satisfied.[16] In a second step we conduct a number of nested and non-nested statistical tests in order to determine which norm gives the most consistent results. Finally, we also use the individual, wage-unit-specific estimates from the RC specification. We look again at the above-mentioned condition concerning the coefficients of expected inflation and reference norms that should also be valid on an individual basis. Furthermore, we will also argue that neither theory nor the actual wage-setting practice in Austria suggest that the individual coefficients should contain a systematic temporal pattern. This fact can also be used to decide about the appropriateness of different reference norms specification. Our overall result is that the leadership norm performs best under all of the employed tests.

3.5 Main results

Tables 4a and 4b present the results of estimating equation (15) with a FE, a RE, a RC and a MG specification. In Table 4a we use the external norm as our measure of reference norms while in Table 4b we use the leadership norm. Furthermore, in both tables we investigate the impact of either GDP growth or the change in unemployment as the measure for real activity.

`Insert Tables 4a and 4b about here`

Looking at Tables 4a and 4b we can make the following observations:

[15] I.e. for 1980–1989, 1990–1999 etc. In robustness tests we have also employed 5 year dummies or a linear time trend. The main results are not affected by these changes.

[16] We write $\hat{\bar{\beta}}_k$ for the estimated cross-sectional average of the individual coefficients $\hat{\beta}_k^i$.

- The differences between the four specifications are rather modest for the estimations with external norms and larger for the ones with the leadership norm. As far as the homogeneous coefficient models are concerned for both specifications the data suggest a FE model. It is important to note, however, that a thorough discussion of the case with homogeneous coefficients is rather superfluous since for both norms the results indicate that there is a considerable degree of **heterogeneity** in the parameter estimations. At the bottom of the tables we report the test statistics for parameter constancy that has been suggested by Swamy (1970) for the RC model.[17] The null hypothesis of homogeneity of coefficients is strongly rejected for all 4 RC specifications in Tables 4a and 4b. Tests based on the individual estimates underlying the MG estimator also show considerable and significant differences between wage-setting units. In light of these results, we will maintain the assumption of heterogeneous coefficients and we will employ the RC model in the following as our benchmark estimation method.

- **Reference norms** are an important factor for the determination of wages. The coefficient for the reference norm is highly statistically significant in all specifications and its size is considerable, typically between 0.5 and 0.6. The theoretical model of section 2 suggests that an importance of reference norms in the neighborhood of $\mu = 0.5$ is sufficient to create reasonable degrees of inflation persistence (cf. Figure 1 for the symmetric case).

- The **expected rate of inflation** over the duration of the contract ($E_{j,\tau}\Delta\tilde{p}_{j,\tau}$), on the other hand, is also an important factor for negotiated wage rates as suggested by New Keynesian theories. Its influence appears to be slightly lower than the influence of reference norms. As described above, the theoretical model underlying equation (15) implies that the (average) estimated coefficients of expected inflation $\hat{\bar{\beta}}_1$ and reference norm $\hat{\bar{\beta}}_3$ should sum up to 1. In Tables 4a and 4b we report the sum of these two coefficients in the lower part of the tables and we also give the statistics and the p-value for the corresponding F-tests. We get a striking result. The null-hypothesis implied by the theoretical model is *never* rejected for the models with heterogeneous coefficients (RC or MG) and the use of leadership norms (columns (3), (4), (7) and (8) in Table 4b), while it is *always* rejected in the remaining cases with external norms (Table 4a) and with the use of homogeneous coefficients (columns

[17] This statistic is distributed χ^2 with $k(P-1)$ degrees of freedom where P is the number of groups (100 in our case) and k is the number of different parameters (6 in our case).

(1) (2), (5) and (6) in Table 4b). These results thus also suggest the use of a model with heterogeneous coefficients and they furthermore give a first strong indication that the leadership norm is probably a more accurate concept than the external measure.

- The **expected development of real activity** contributes to the size of the wage increase. If GDP is expected to grow faster by 1% this will increase the average wage agreement by between 0.08% and 0.33%. On the other hand, if unemployment rate is expected to decrease by 1 percentage point this is expected to boost wage claims by between 0.22% and 1.23%. Interestingly, these different values for GDP growth and the change in unemployment imply an Okun coefficient of about 4 which is broadly compatible with empirical estimations for Austria. An interesting result that is crucial for our argument is related to the fact that the coefficient of expected inflation is lower and the coefficient of expected real activity much lower for the leadership norm. This result is directly compatible with the wage leadership story where it can be assumed that the wage-leaders in the metal sector look closely and thoroughly at the macroeconomic variables. The other wage-setters that follow, however, orient themselves primarily on the result on the metal sector. Partly, because they lack the resources to engage in time-consuming forecasts and partly because they trust the accurateness of the general assessment in the wage-leading sector. As a result, the broad macroeconomic outlook is already contained in the leadership norm and over and above this norm there is little influence on the negotiated wages.

3.6 Nested and non-nested tests on the importance of reference norms

In Tables 4a and 4b we have only focused on two commonly used reference norms. In this section we will also look at the two other norms that correspond to proposals in the related literature: the habit norm and the price indexation norm. We want to use formal statistical tests to investigate in a more systematic way which one of the four norms is the most appropriate concept. It is, however, quite difficult to distinguish between the alternative formulations since most of them are highly correlated. It is therefore not surprising that if we estimate the benchmark equations in Table 4a and 4b with the two alternative norms (results not shown) they also come out with a significant positive coefficient. It is, however, important to note that also for these alternative concepts (as was the case for the external norm) the condition that $\hat{\bar{\beta}}_1 + \hat{\bar{\beta}}_3 = 1$ can be rejected for all specifications.

In order to get an idea about the relative importance of the different norms we first use non-nested J-tests (cf. Smith, 1996; Greene, 2003, chap. 8): We regress the dependent variable $\Delta w^i_{j,\tau}$ on the benchmark set of regressors including the change in reference norm $\Delta rn^{1,i}_{j,\tau}$. The fitted values of this regression are then included in a regression of $\Delta w^i_{j,\tau}$ on the benchmark variables plus an alternative norm $\Delta rn^{2,i}_{j,\tau}$. If Δrn^2 is the correct measure then the coefficient on the fitted values from the first regression should be close to zero (which is determined by a t-test). In a next step we reverse the roles of Δrn^1 and Δrn^2. Unfortunately, this test does not guarantee unambiguous results since it is possible that we reject an independent role of Δrn^1 in the first and of Δrn^2 in the second regression. In Table 5 we report the results of the J-tests. We follow the practice of Smith (1996) and interpret the relative size of the t-statistics as an indicator of "dominance" in order to deal with the inconclusive cases. Overall, the results in Table 5 suggest that the leadership norm is the most consistent and important measure of reference norms. It is the dominant measure in all pairwise comparisons and in the case of the price indexation norm the alternative measure is not even statistically significant. The external norm is also an influential measure that is dominant in all regressions except the one where it is paired with the leadership norm.

`Insert Table 5 about here`

Similar conclusions also result from nested tests in which all pairwise combinations of reference norms are included at the same time into the benchmark estimation. This is shown in Table 6. Although collinearity is likely to affect the estimates, the coefficient on the leadership norm always stays significant and also its variation in size is rather small. The external norm, on the other hand, becomes much smaller (although still statistically significant) once entered together with the leadership norm while it seems to be rather stable in the other comparisons. The coefficients on the habit norm and the price indexation norm are rather small and often insignificant.

`Insert Table 6 about here`

On balance, these results suggest that the leadership norm seems to be the most appropriate description of the reference norm that influences the process of wage-setting in Austria. This conforms with the results based on the condition that $\hat{\bar{\beta}}_1 + \hat{\bar{\beta}}_3 = 1$.

3.7 An analysis of the individual heterogeneous coefficients

So far we have focused our analysis on the average values of the empirical models with heterogeneous coefficients. For our purpose it is, however, also useful to look at the individual coefficients that underlie these common measures. An analysis along these lines can deliver important insights into the nature of reference norms and the temporal and possibly asymmetric structure of wage-setting.

The first step in this context is to look again at the implication of the theoretical model. The condition that the coefficients of the reference norm and of expected inflation should sum up to 1 is supposed to hold not only in the aggregate but also for each of the 100 wage-setting units individually. We have conducted F-tests using the results from the individual estimations that underlie the RC specification. In Table 7 we report the percentage of wage-setting units for which the individual F-tests imply a rejection of the condition for various assumptions about the nature of reference norms. For the case of leadership norms the theoretical implication is rejected for 15 of the 100 wage-setting units (using a 5% level of significance; for the 1% level the condition is violated in only 6% of the cases). The rejection rate is much higher for the external, the habit and the price indexation norm where it comes out as 69%, 89% and 27%, respectively. Also for the case where we abstract from reference norms, the individual coefficients of expected inflation (which is now the only nominal variable on the right-hand-side of the wage-setting equation) are in 23% of the units significantly larger than unity.[18] The F-tests based on the individual coefficient estimations thus again confirm the conclusions based on the average coefficients and on the nested and non-nested tests.

`Insert Table 7 about here`

The individual coefficients can, however, also be used to provide further evidence on the appropriateness of the different reference norms. This argument is based on an analysis of the temporal pattern of the reaction to the macroeconomic variables and the reference norm. In particular, we can order the individual regression coefficients $\hat{\beta}_1^i$ to $\hat{\beta}_3^i$ with respect to the (median) month in which each of the 100 wage-setting units typically concludes its wage agreements. A-priori there exists no reason to believe that there should be any discernible temporal pattern in the size of these coefficients (e.g. that they should be higher in spring than in fall or vice versa). In fact, if the assumption of complete

[18]Note that in this case the F-test on the common-mean coefficient of expected inflation also rejects the null-hypothesis of a value of 1.

symmetry in wage-setting is valid (as assumed in the model of section 2.2) then the coefficients should be identical across wage-setting units. All possible differences in the estimated coefficients should then only be due to measurement or random errors and thus should not show any noticeable temporal structure. In Figure 2 we present pictures that allow to analyze this topic in more detail for the case without norms (panel A), with external norms (panel B) and with leadership norms (panel C). For each norm we show the size of $\hat{\beta}_3^i$ and of $\hat{\beta}_2^i$.[19]

`Insert Figure 2 about here`

We start our discussion with the case that abstracts from reference norms and that corresponds to the standard wage-setting equation in New-Keynesian models. The pictures in panel A of Figure 2 give an unambiguous and striking result. There seems to be a clear temporal pattern in $\hat{\beta}_2^i$, i.e. in the reaction to expected changes in the unemployment rate. It is strongest for the sector that negotiates in November and afterwards decreases in importance. In fact, the average coefficient for the units contracting in November (-1.84) is more than double the size than for the rest of the economy (-0.74). This is in itself a strong indication that the November agreements play a special role in the system of wage-setting in Austria and that there seem to be noticeable asymmetries across sectors.[20] The wage leading units apparently put a considerably higher weight on the general macroeconomic situation as reflected by their expectations about future unemployment.

If we look at the distribution of the individual coefficients when external norms are used (panel B of Figure 2) we find a similar picture. In this case the RC model reveals again a clear temporal pattern where the importance of reference norms seems to increase over the year starting in November while the (absolute values of the) coefficients of unemployment decrease from November to the summer of next year. The assumption of symmetric reference norms is thus again contradicted by the data.

If we repeat the same exercise for the leadership norm we find different results as shown in panel C of Figure 2. There is no discernible temporal pattern in the importance of

[19] The individual coefficients shown in the figures are the ones from the RC estimation. Here we had to deal with an issue that arises in the empirical estimations. Since the leadership norm is zero for the four wage-leading units their values are dropped from the RC estimation altogether and we would not have results for these crucial units. In order to prevent this from happening we have set for these 4 units the leadership norm equal to 1 in the year 1980. This amounts to an additional year dummy for these four units that is, however, not statistically significant and leaves all coefficients for the other units virtually unchanged. Moreover, the use of the individual OLS regressions that underlie the mean group estimator leads to completely parallel results.

[20] There are 14 units that typically contract in November, from which we have categorized 4 as the wage leaders. The average coefficient for the wage-leaders is -2.08.

reference norms anymore. Their average value is almost identical for the sectors following the agreements in November. The same is also true for the estimated reaction to expected changes in unemployment ($\hat{\beta}_2^i$) where the temporal pattern is weak (or basically non-existing). The four units that comprise our group of wage leaders show the highest reaction to changes in the unemployment rate (between -1.99 and -2.2) while the other 10 units that contract in November reveal a reaction that is more similar to the rest of the economy except two units (white collar employees in the chemical and the paper industry for which the estimated coefficient is around -1.3). This suggests that these other units already have information about the negotiated wage in the metal sector and use this as a guideline for their own wage agreements.

The results taken together imply that the assumption of symmetric reference norms is not appropriate for the Austrian situation. One possible explanation for the observed temporal patterns in the case of external or non-existing norms could be that wage-setting units are in fact ordered over the year with respect to the strength with which they react to the macroeconomic situation. An inspection of the nature of the 100 units does not support this hypothesis, however, since the units that contract later in the year include a number of sectors that are typically assumed to react rather strongly to business cycle conditions (like, e.g., the construction sector which contracts in May). Also the increasing temporal pattern of the coefficients of reference norms $\hat{\beta}_3^i$ is rather implausible and contradicts anecdotal evidence that — if anything — the importance of norms should *decrease* over the course of a wage round (see Traxler et al., 2008). On the whole, we would thus argue that it is more suggesting to interpret the observable temporal patterns in the estimations that use no or external reference norms as an indication of misspecifications of the empirical relation.

The results of this section therefore further support our overall conclusion that wage-setting in Austria is characterized by a system of wage leadership by the metal sector. In the negotiations of fall ("autumn bargaining round") the wage leader mostly focuses on the macroeconomic conditions that are expected to prevail in the upcoming year and puts no (or at least much less) weight on the wage rates that have been set in the negotiation rounds that ended before summer. This is different for the wage agreements following the metal sector that seem to regard these earlier settlements as a benchmark that acts as a reference point for the success of their own wage agreements.[21] The fact that the

[21]The results on wage leadership confirm parallel findings by Traxler et al. (2008) who work with a similar dataset on collectively bargained wages in Austria and also conclude that the metal sector acts as the wage leader. Their approach to the topic is, however, more from the perspective of industrial

wage leader is mostly concerned with the general macroeconomic environment and less with wage comparisons is likely to contribute to a system that shows less endogenous persistence than in other countries.

3.8 Further Robustness Tests

So far we have used a rather parsimonious specification to study the role of reference norms in wage-setting. In this section we want to check the robustness of our benchmark estimations with leadership norms by using different samples and including additional regressors. In Table 8 we present the results when the expected change in the unemployment rate is used as the measure for real activity. The results (not shown) for GDP growth as a measure of real activity are similar.

`Insert Table 8 about here`

In columns (2) and (3) of Table 8 we look at the benchmark estimation for two different time samples (before and after 1993). The results are fairly similar although the coefficient of the reference norm is smaller and the one of expected inflation is higher for the later sample. This would suggest that after the middle of the 90ies the rate of inflation has become a more important anchor for wage-setting than it has been the case before. One would speculate that this might be at least partly connected to the influence of European monetary integration and to the "Great Moderation" in general. If we focus only on the private sector (79 instead of 100 wage-setting units) the main results are also unchanged (see column (4)). As expected, the private sector seems to react more sensitively to business cycle conditions although the difference is not statistically significant. In column (5) we have pooled similar wage-setting units together (i.e. units in identical sectors that negotiate at the same time of the year and reach very similar agreements). This leaves us with 55 "independent" instead of the 100 total units. Using these independent units strengthens the role of wage leadership since the coefficient of the reference norm increases (to 0.63) and the coefficient of the change in unemployment decreases and becomes insignificant.

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relations and they provide more detailed evidence on the emergence of this system around 1980 and on the specific role played by employer organisations and unions. In contrast to their work, we derive the estimation equation from an explicit theoretical model, we focus on the implications of asymmetries in reference norms on inflation persistence, we use explicit forecasts for the macroeconomic variables and we also consider alternative hypotheses concerning reference norms.

In column (6) we include forecast errors as suggested by the theory-based formulation in (15). This considerably increases the role of expected real activity but leaves the coefficients of all other variables basically unchanged. Also the coefficients of the forecast errors themselves are not statistically significant which can be regarded as a confirmation of the assumption of rational expectations.

In column (7) we include lagged inflation in the benchmark equation. This corresponds to a "hybrid" Phillips curve that is quite popular in the recent macroeconomic literature. The results are almost unchanged and the coefficient of past inflation is almost zero, thereby confirming that the reference norm is not only a proxy for past inflation. One would expect that a longer duration of a wage contract is associated with higher wage increases. This is in fact borne out by the data as shown in column (8) of the table although the effect is rather small.

If we add the expected *level* of unemployment instead of the change in unemployment as a measure of real activity (column (9)) this slightly increases the coefficients on reference norms while the coefficient of the level of unemployment itself is insignificant. It is interesting to note, however, that if we include the expected and the lagged level of unemployment together (column (10)) then these two coefficients are of approximately equal size but have different signs (as confirmed by a F-test). This implies that the expected change in unemployment is in fact the more appropriate specification as is suggested by our theoretical model that implies a wage curve.

We have also conducted a number of other robustness checks that are not shown in Table 8. In these further specifications we have included, e.g., monthly dummies, additional lagged variables, sectoral unemployment rates. Furthermore we have also conducted a number of estimations where we have instrumented the explanatory variables of the benchmark equation in order to account for possible endogeneity. The main results have again been unaffected by these changes. Finally, we have also run the benchmark regression specification while assuming that other sectors are the wage leaders. This led to unconvincing results.

The basic message of the robustness tests in Table 8 is that the main results of the benchmark specifications are unchanged: reference norms have a considerable impact on collective wage agreements and their weight seems to be at least as high as the one of expected inflation. This holds for a large number of samples, different specifications and in the presence of various additional variables. Furthermore, as shown at the bottom of Table 8, the condition that $\hat{\bar{\beta}}_1 + \hat{\bar{\beta}}_3 = 1$ is only rejected for one of the alternative specification (column (10)) at the 5% level and for none at the 1% level.

4 Conclusions

In this paper we have studied the influence of reference norms in an otherwise standard Taylor model with staggered wages. We have shown that the inclusion of reference norms considerably increases inflation persistence. On the other hand, we have also shown that the impact on persistence very much depends on the precise definition of reference norms and on possible asymmetries in their importance between sectors. In the empirical section we have documented that reference norms play an important and significant role in the setting of collectively bargained wages in Austria. The impact of our most preferred measures of reference norms is typically larger than the impact of expected inflation and in our benchmark specification its weight is about 60%. The theorectical model of section 2 suggests that a weight of this magnitude is sufficient to produce a reasonable degree of inflation persistence. Comparable weights of backward indexation in standard price-setting models look much more implausible and at odds with the existing survey evidence. When we compare different concepts of reference norms we find that the leadership norm gives the best description of the data and that it leads to superior results than the assumption of external, habit or price indexation norms. These results are confirmed by tests that use theory-based restrictions on the average parameter estimates, by explicit nested and non-nested statistical tests and also by an analysis of the temporal pattern of the individual estimated coefficients. All specifications support the hypothesis that wage leadership by the metal sector plays a crucial role for wage-setting in Austria.

Taken together, the theoretical and empirical results suggest that differences in reference norms or the existence of wage leadership can be at least partly responsible for the observed cross-country differences in inflation persistence and in wage rigidity (cf. Cecchetti and Debelle, 2006; Dickens et al., 2007; Holden and Wulfsberg, 2007). Different reference norms and/or intersectoral differences in the importance of norms can be reflected in the dispersion of the aggregate measures of persistence across countries. In this context it is, e.g., interesting to note that the results in Cecchetti and Debelle (2006, Table 2) indicate that Austria is one of the countries with the lowest degree of inflation persistence (rank 17 among 19 industrialized countries). This could at least partly reflect the influence of wage leadership.[22] By the same token, our results also offer an explanation

[22] It would be interesting to study the role of reference norms also for other countries. In order to construct a similar dataset as the one used in this paper one would either have to use (official) data on collectively bargained wages (as for Austria) or directly resort to wage registers maintained by wage setters (unions or employer federations). This should be possible for a number of European countries (e.g. Germany or the Scandinavian countries).

for the often weak correlation between various measures of wage-setting institutions and the observed degrees of wage rigidity. A recent summary paper concludes, e.g., that “the connection between unions and wage rigidity, although it may seem obvious in theory, appears somewhat shakier in our data than one might expect” (Dickens et al., 2007, p. 211). The results of our paper imply that the shaky and often nor very robust findings concerning the relation between inflation persistence or real wage rigidity and labor market institutions are probably caused by the fact that these institutional variables do not include measures for the importance, the prevalent type and for possible asymmetries of reference norms.

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Appendices

A The microfounded model

A.1 Production

The set-up of the model follows Ascari (2000, 2003) and we only want to sketch it here rather briefly. There is a continuum of industries $i \in [0, 1]$ and two sectors (A and B) of equal size, where sector A consists of the industries in $i \in [0, \frac{1}{2}]$ and sector B of those in $i \in [\frac{1}{2}, 1]$. The wages in both sectors are set by unions where we assume that there is one union that is attached to each firm. Furthermore, we assume that workers are attached to their sector and there is no labor mobility between sectors.[23] Wages are fixed for two periods and sector A unions set their wages in periods $t = 0, 2, 4, ...$ while unions in sector B decide in periods $t = 1, 3, 5, ...$.

There is a homogeneous output good Y_t that is produced by competitive firms with the following CES production function:

$$Y_t = \left(\int_0^1 Y_{it}^{\frac{\theta - 1}{\theta}} di \right)^{\frac{\theta}{\theta - 1}}, \tag{16}$$

where the Y_{it} are intermediate inputs that are necessary to produce the final output and $\theta > 1$ is the elasticity of substitution between these different intermediate goods. This leads to the demand functions for intermediate goods:

$$Y_{it} = \left(\frac{P_{it}}{P_t} \right)^{-\theta} Y_t, \tag{17}$$

where the aggregate price index is given by:

$$P_t = \left(\int_0^1 P_{it}^{1-\theta} di \right)^{\frac{1}{1-\theta}} \tag{18}$$

The intermediate goods Y_{it} are produced by the firms indexed by $i \in [0, 1]$. Firms

[23]This is a crucial assumption as argued by Ascari (2003): "*Only models with some form of labour immobility could potentially deliver a substantial degree of persistence*" (p. 527, original emphasis).

have access to a production function:

$$Y_{it} = A_t L_{it}^{\alpha}, \tag{19}$$

where $0 < \alpha \leq 1$ and where L_{it} is the amount of labor used by firm i in period t. For the purpose of this paper we set $A_t = 1$ and $\alpha = 1$ (constant returns to scale). The firms are assumed to be price takers (perfect competition in intermediate goods production). They thus set prices equal to marginal costs or:

$$P_{it} = W_{it} \tag{20}$$

where W_{it} is the wage rate that firm i faces in period t. Put differently, firms will hire labor until the real wage equals the marginal product of labor which in this case is just $A_t = 1$.

We can insert (17) and (19) into (20) in order to derive an expression for labor demand L_{it} in terms of the wage rate W_{it}.

$$L_{it} = W_{it}^{-\theta} \left(P_t^{\theta} Y_t \right), \tag{21}$$

A wage-setting union uses equation (21) to take into account the effect of an increase in the wage rate W_{it} on labor demand L_{it}. Due to the assumption that unions are atomistic they neglect any possible effect of their wages on the aggregate variables P_t and Y_t.

A.2 Households

The intertemporal utility function is given by:[24]

$$U_{j0} = \sum_{s=0}^{\infty} \beta^s u(C_{js}, \frac{M_{js}}{P_s}, L_{js}) \tag{22}$$

where β is the time discount factor, C_{js} is real consumption, $\frac{M_{js}}{P_s}$ are real money holdings and L_{js} is labor supply by household j in period s. There is a series of budget constraints:

$$P_t C_{jt} + M_{jt} + B_{jt} = \xi_t M_{jt-1} + (1 + i_{t-1}) B_{jt-1} + W_{jt} L_{jt} + T_{jt} + H_{jt} \tag{23}$$

[24]We abstract in the following from uncertainty as in Ascari (2000). As noted there (FN 7) the introduction of uncertainty would be straightforward.

The nominal income in period t consists of a predetermined level of wealth, given by the money balances M_{jt-1} and the amount and interest earned on bonds that are carried over from period $t-1$, i.e. $(1+i_{t-1})B_{jt-1}$. Money holdings are subject to a common multiplicative shock ξ_t (see Ascari, 2003). In addition households have labor income $W_{jt}L_{jt}$ and a lump-sum government transfer T_{jt}. Households might also receive insurance payments H_{jt} that occur in the presence of monetary shock.[25] The total nominal income can be used for purchases of consumption P_tC_{jt}, money M_{jt} and bonds B_{jt}.

Maximization of (22) with respect to C_{jt}, M_{jt} and B_{jt+s} leads to the FOCs:

$$u_C(C_{jt}, \frac{M_{jt}}{P_t}, L_{jt}) = \beta(1+i_t)\frac{P_t}{P_{t+1}}u_C(C_{jt+1}, \frac{M_{jt+1}}{P_{t+1}}, L_{jt+1}) \tag{24}$$

$$u_{\frac{M}{P}}(C_{jt}, \frac{M_{jt}}{P_t}, L_{jt}) = u_C(C_{jt}, \frac{M_{jt}}{P_t}, L_{jt})\left(\frac{i_t}{1+i_t}\right), \tag{25}$$

where $u_x(C_{jt}, \frac{M_{jt}}{P_t}, L_{jt}) = \frac{\partial u(C_{jt}, \frac{M_{jt}}{P_t}, L_{jt})}{\partial x}$. Equations (24) and (25) represent the Euler equation for consumption and the money demand equation. For a more compact expression we will write in the following: $u_x(C_{jt}, \frac{M_{jt}}{P_t}, L_{jt}) = u_x(t)$.

Union j sets the wage for two periods, i.e. under the constraint that $W_{jt} = W_{jt+1}$. The unions take into account that labor demand L_{jt} by the firm is given by (21) and that the income of the household $W_{jt}L_{jt}$ also depends on this magnitude. In particular: $L_{jt} = W_{jt}^{-\theta}P_t^{\theta}Y_t$, $W_{jt}L_{jt} = W_{jt}^{1-\theta}P_t^{\theta}Y_t$ (and similar for the second period). Maximization of (22) with respect to W_{jt} taking these relations into account thus leads to the wage-setting equation:

$$W_{jt} = -\frac{\theta}{\theta-1}\frac{u_L(t)P_t^{\theta}Y_t + \beta u_L(t+1)P_{t+1}^{\theta}Y_{t+1}}{u_C(t)\frac{P_t^{\theta}Y_t}{P_t} + \beta u_C(t+1)\frac{P_{t+1}^{\theta}Y_{t+1}}{P_{t+1}}} \tag{26}$$

This corresponds to equation (15) in Ascari (2000) and to equation (22) in Huang and Liu (2002) in a multiperiod framework. In the absence of staggering (26) reduces to the usual optimality condition: $\frac{W_{jt}}{P_t} = -\frac{\theta}{\theta-1}\frac{u_L(t)}{u_C(t)}$.

[25]On this see Ascari (2000, 670) and Huang and Liu (2002).

A.3 Linearizations

We can linearize the FOCs of this model around a zero inflation steady state. The linearized wage-setting equation can be written as:

$$w_{jt} = bp_t + (1-b)p_{t+1} + \gamma(by_t + (1-b)y_{t+1}), \tag{27}$$

where lower case letters stand for deviations around the steady state. Furthermore, $b = \frac{1}{1+\beta}$ and $\gamma = \frac{-\eta_{cc}+\eta_{ll}}{1+\theta\eta_{ll}}$ where $\eta_{xx} \equiv \frac{u_{xx}x}{u_x}$ is the elasticity of the marginal utility of x with respect to x. In particular, η_{cc} and η_{ll} are the inverses of the intertemporal elasticity of substitution of consumption and of labor supply, respectively. Note that for the general case with decreasing returns to scale ($\alpha \leq 1$) we would get: $\gamma = \frac{-\eta_{cc}+\frac{1}{\theta(1-\alpha)+\alpha}\eta_{ll}}{1+\frac{\theta}{\theta(1-\alpha)+\alpha}\eta_{ll}}$.[26] The wage w_{jt} stands for the wage that is set in period t by union j. Since all unions in a sector are assumed to be identical we can write $w_{jt} = w_t^A$ and $w_{jt} = w_t^B$ if union j belongs to sector A (B). Adding the expectation operator and allowing for a sector specific γ^i equation (27) corresponds to (2) in the main text.

The price index (18) for the two-sector model is given by

$$P_t = W_t = \left(\frac{1}{2}\left(W_t^A\right)^{1-\theta} + \frac{1}{2}\left(W_t^B\right)^{1-\theta}\right)^{\frac{1}{1-\theta}}$$

Linearizing this around the steady state leads to:

$$p_t = \frac{1}{2}\left(w_t^A + w_t^B\right) \tag{28}$$

This corresponds to equations (4) in the paper.

Finally we can also linearize the FOC conditions (24) and (25). It can be shown that under specific assumption concerning the multiplicative shock on money holdings ξ_t the velocity of money is constant over time. In this case the linearization of (25) leads to:[27]

$$-\eta_{cc}y_t = m_t - p_t \tag{29}$$

In the paper we focus on the case where $\eta_{cc} = -1$ (which corresponds to a utility function that is logarithmic in consumption). This is stated in equation (5).

[26] See Ascari (2000, p. 674) and Ascari (2003, 520f.).
[27] See Ascari (2003, p. 514f.) and Ascari (2000, FN 23).

B The derivation of λ^i

We can insert (4), (5) and (6) into (1) to derive:

$$w_t^i = \psi_1^i w_{t-1}^{-i} + \psi_2^i E_t w_{t+1}^{-i} + \psi_3^i m_t \tag{30}$$

where $\psi_1^i = \frac{b(1-\gamma^i)(1-\mu^i)+2\mu^i}{1+\mu^i+\gamma^i(1-\mu^i)}$, $\psi_2^i = \frac{(1-b)(1-\gamma^i)(1-\mu^i)}{1+\mu^i+\gamma^i(1-\mu^i)}$, $\psi_3^i = \frac{2\gamma^i\left(1-\mu^i\right)(b+\rho(1-b))}{1+\mu^i+\gamma^i(1-\mu^i)}$. Note that for $t = 0, 2, 4, ...$ it holds that $i = A$ and $(-i) = B$, while for $t = 1, 3, 5, ...$ we have that $i = B$ and $(-i) = A$. There exists various ways to solve this forward-looking difference equation. We choose the method of undetermined coefficients and start with the conjecture that the solution for w_t^A is of the form:

$$w_t^A = \lambda^A w_{t-1}^B + \theta^A m_t \tag{31}$$

The parallel conjecture for w_{t+1}^B is that:

$$w_{t+1}^B = \lambda^B w_t^A + \theta^B m_{t+1} \tag{32}$$

From (32) together with (6) it follows that $E_t w_{t+1}^B = \lambda^B w_t^A + \theta^B \rho m_t$. We can insert this expression into (30) to derive that:

$$w_t^A = \frac{\psi_1^A}{1-\psi_2^A\lambda^B} w_{t-1}^B + \frac{\psi_2^A\theta^B\rho + \psi_3^A}{1-\psi_2^A\lambda^B} m_t \tag{33}$$

In an analogous way we can derive that:

$$w_{t+1}^B = \frac{\psi_1^B}{1-\psi_2^B\lambda^A} w_t^A + \frac{\psi_2^B\theta^A\rho + \psi_3^B}{1-\psi_2^B\lambda^A} m_{t+1} \tag{34}$$

Comparing (31) with (33) and (32) with (34) we see that the equilibrium values for λ^A, λ^B, θ^A and θ^B are implicitly given as the solutions to the following system of four equations: $\lambda^A = \frac{\psi_1^A}{1-\psi_2^A\lambda^B}$, $\lambda^B = \frac{\psi_1^B}{1-\psi_2^B\lambda^A}$, $\theta^A = \frac{\psi_2^A\theta^B\rho+\psi_3^A}{1-\psi_2^A\lambda^B}$ and $\theta^B = \frac{\psi_2^B\theta^A\rho+\psi_3^B}{1-\psi_2^B\lambda^A}$. For the crucial parameter λ^i the solutions come out as:

$$\lambda^A = \frac{2\psi_1^A}{1-\psi_2^A\psi_1^B+\psi_1^A\psi_2^B+\sqrt{(1+\psi_2^A\psi_1^B-\psi_1^A\psi_2^B)^2-4\psi_2^A\psi_1^B}} \tag{35}$$

$$\lambda^B = \frac{1 + \psi_2^A \psi_1^B - \psi_1^A \psi_2^B - \sqrt{\left(1 + \psi_2^A \psi_1^B - \psi_1^A \psi_2^B\right)^2 - 4\psi_2^A \psi_1^B}}{2\psi_2^A} \tag{36}$$

For the symmetric case with $\mu^A = \mu^B = \mu$ and $\gamma^A = \gamma^B = \gamma$ (and in addition with $b = \frac{1}{2}$) we can derive that:

$$\lambda^A = \lambda^B = \lambda = \frac{1 - \sqrt{1 - 4\psi_1\psi_2}}{2\psi_2}$$

where $\psi_1 = \frac{1}{2}\frac{(1-\gamma)(1-\mu)+2\mu}{1+\mu+\gamma(1-\mu)}$ and $\psi_2 = \frac{1}{2}\frac{(1-\gamma)(1-\mu)}{1+\mu+\gamma(1-\mu)}$. Inserting these values for ψ_1 and ψ_2 gives (9) in the main text. We can take this expression to calculate:

$\frac{\partial\lambda}{\partial\gamma} = -\frac{1+\mu^2+\gamma(1-\mu^2)-2\sqrt{\gamma(1-\mu^2)+\mu^2}}{(1-\gamma)^2(1-\mu)\sqrt{\gamma(1-\mu^2)+\mu^2}} < 0$

$\frac{\partial\lambda}{\partial\mu} = -\frac{2\left(\gamma(1-\mu)+\mu-\sqrt{\gamma(1-\mu^2)+\mu^2}\right)}{(1-\gamma)(1-\mu)^2\sqrt{\gamma(1-\mu^2)+\mu^2}} > 0$

Note that from (31) and (32) we can derive that:

$$w_t^A = \lambda^A w_{t-1}^B + \theta^A m_t = \Lambda w_{t-2}^A + \theta^A m_t + \lambda^A \theta^B m_{t-1}$$

$$w_{t+1}^B = \lambda^B w_t^A + \theta^B m_{t+1} = \Lambda w_{t-1}^B + \theta^B m_{t+1} + \lambda^B \theta^A m_t,$$

where we have defined that $\Lambda \equiv \lambda^A \lambda^B$. This measure for annual persistence is the same in both sectors even if $\mu^A \neq \mu^B$ and $\gamma^A \neq \gamma^B$. For this reason we use it frequently as our preferred measure of persistence in the main text.

C The data

C.1 Sectoral wage change data

C.1.1 The basic index series

Our collective bargaining wage data are based on detailed series of the *Tariflohnindex* (TLI, "Index of Agreed Minimum Wages"), a monthly database maintained by Statistics Austria. The TLI is a Laspeyres index of sectoral collectively negotiated minimum wages according to a particular base year. Predominantly, our data are from the TLI 1986 (for the years 1986 - 2006). For 1979 - 1985 we use the TLI 1976. Individual series are chained with the ratio of the annual average TLI 1986 to the TLI 1976 in 1986.[28] Wage changes in a particular year are the relative changes of the TLI index compared to the last change. In total, we thus have 27 years of collective wage changes. As the system of wage leadership is only in place since the beginning of the 1980s (cf. Traxler et al., 2008) we do not use data from previous years.

Typically, these indices are analyzed only at rather high levels of aggregation. However, Statistics Austria also published more detailed monthly index values which are - at the most disaggregate level - identical, or come close to, individual collective agreements. There are 125 such indices in the TLI 1986 which are mostly for branches and separately for blue-collar and white-collar workers. While most of these refer to collective agreements for the whole country, some of these series cover only one federal state *(Bundesland)*. Sometimes, there is also a distinction between "industry" and "trade" *(Gewerbe)*. All of this reflects the practice of collective bargaining in Austria. Some examples are "blue-collar workers in the paper-processing trades", "blue-collar workers in the metal industry", "white-collar workers in the clothing and textile industry in Vorarlberg", and "white-collar workers in insurance companies".

Furthermore, there are several indices for the public sector (e.g. for workers in the federal administration) and a few series for (mainly) publicly owned transport companies (such as the *Österreichische Bundesbahnen*). A full list of all 125 TLI series is given in Table C-1 which is structured along the partiton of collective agreements in blue-collar, white-collar and public sector workers as well as workers in public transport. Within each

[28] For detailed information we refer to Statistics Austria publications such as *Tariflohnindex 1986. Aufbau und Gewichtung* (Vienna, 1988) and *Tariflohnindex 1976. Aufbau und Gewichtung* (Vienna, 1978). Basic information can also be found in Statistics Austria's monthly publication *Statistische Nachrichten* in no. 6/1978 and no. 1/1988. We thank Markus Bönisch and Helga Maurer for delivering these data and giving us background information.

partition, the series names are listed in decending order of the TLI 1986 weight.

Table C-1: List of all 125 basic TLI 1986 units

unit	quality	unit	quality	unit	quality
Blue-collar workers:		*White-collar workers:*		*Public sector:*	
A_I_Metall	1	N_H_Handel	1	O_B_Admin	2
A_G_Metall	1	N_I_Metall	1	O_L_Lehrer	1
A_F_Gastro	1	N_G_Allgemein	1	O_L_Admin	1
A_G_Bau	1	N_B_Versicherung	1	O_B_Post	1
A_H_Handel	1	N_B_Banken	1	O_B_Lehrer	1
A_G_Bauhilf	1	N_I_Chemie	1	O_L_AdminWien	1
A_G_Holz	1	N_G_Bau	1	O_B_Wache	1
A_I_Chemie	1	N_B_Sparkassa	1	O_S_SV	1
A_I_Nahrung	3	N_I_Ewerk	1	O_G_Mittel	1
A_I_Bau	1	N_I_Bergwerk	1	O_G_Gross	1
A_I_Textil	2	N_I_Nahrung	3	O_G_Klein	1
A_G_Sonstige	3	N_B_Raiffeisen	1	O_B_Uni	1
A_G_Druck	1	N_R_Ziviltechnik	3	O_L_LehrerWien	1
A_I_Holz	1	N_G_Druck	1	O_S_SP	3
A_I_Stein	1	N_F_Gastro	1	O_S_Kirche	3
A_V_Gueter	2	N_I_Bau	1	O_B_Heer	1
A_G_Friseur	3	N_V_Spedition	1	O_B_Richter	1
A_G_Fleischer	1	N_R_Steuerberater	3	O_B_Theater	1
A_I_Bekleidung	3	N_I_Stein	1	O_S_OEGB	1
A_G_Baecker	1	N_I_TextilVbg	1		
A_I_Ewerk	2	N_B_Volksbank	1	*Public transport:*	
A_I_Papier	2	N_V_ORF	1	V_V_OEBB	1
A_L_Gut	1	N_L_Lager	1	V_V_Schienen	3
A_G_Nahrung	3	N_R_Apotheke	3	V_V_WVB	1
A_I_Bergwerk	1	N_I_Holz	1		
A_G_Textil	3	N_I_Textil	1		
A_I_Saege	1	N_I_Erdoel	1		
A_V_Bus	1	N_V_Luft	3		
A_I_Lederver	3	N_G_Werbung	1		
A_L_Forst	1	N_R_Arzt	3		
A_L_Lager	1	N_G_Fleischer	1		
A_I_Glas	2	N_I_Bekleidung	2		
A_I_Erdoel	1	N_F_Fachverband	3		
A_I_Pappe	1	N_I_Papier	1		
A_V_Spedition	1	N_F_Kur	3		
A_G_Plastik	1	N_F_Reisebuero	1		
A_F_Kur	2	N_L_Gut	1		
A_G_Putzerei	2	N_G_Baecker	2		
A_G_Leder	3	N_R_Anwalt	3		
A_G_Chemie	1	N_R_Zahnarzt	3		
A_V_Tank	1	N_I_Pappe	1		
A_I_Gas	2	N_I_Gas	1		
A_V_Luft	1	N_I_Schuh	1		
A_G_Papier	1	N_I_Glas	1		
A_F_Fachverband	3	N_I_Saege	1		
A_V_Schiff	3	N_G_Kaese	2		
A_I_Lederprod	2	N_F_Amuesier	1		
A_F_Kino	1	N_V_Fahrschule	2		
A_V_Schienen	1	N_B_PSK	1		
		N_I_Film	1		
		N_L_Forst	1		
		N_V_Schiff	3		
		N_R_Notar	3		
		N_V_Schienen	1		

Altogether, these 125 basic series of the Tariflohnindex - or as we call them - "TLI units" or simply "units" represent 194 collective agreements.[29] Mostly, a TLI index represents only one collective agreement. However, there are cases when such an index covers two ore more agreements. The most notable examples are the series for blue-collar workers in the food industry which aggregates 10 agreements, the series for white-collar workers in wholesale and retail trade representing 10 agreements and the series for blue-collar workers in "supplementary construction trades" *(Bauhilfsgewerbe)* standing for 9 agreements.

C.1.2 Data checks and "quality" of the TLI series

Previous knowledge and the observation of collective bargaining in Austria suggest that most collective agreements last for exactly one year. The question was whether such a regular pattern could also be observed in the data.

Each of the 125 TLI series was scrutinized in which months there was a change in the index and by how much it changed. All series are monotonically increasing in time (collective wages are nominally rigid downwards). Besides that, there was a substantial amount of irregularities. For example, the month of the wage change was not constant. It shifted especially in the 1970s where - apparently - the time structure of wage negotiations shifted markedly (putting the system of wage leadership into place). Moreover, in quite a lot of units there was more than one change of the index in one year whereas in other series or years there was no change at all (most notably in the public sector where there was a "wage freeze" in 1996 and 1997). In such cases we tried to distinguish with great care between "main" and "minor" changes in the index. (Some of these minor changes were easily explained by rounding differences due to chaining the TLI 1986 with the TLI1976, to the introduction of the Euro or to changes in data processing.) In most cases, it was quite obvious to determine *the* month of the wage change. With this information we computed the associated wage change between those months (where it was possible that there was no wage change at all in a particular year).

There were, however, also units where the judgement was difficult, e.g. where there were many index changes spread over the year. Based on the overall impression of how regularly (i.e. not too often or not too rarely) the index figures increased we subjectively

[29] There are considerably more such agreements, however our data comprise all major collective agreements and the data are constructed to give a representative picture of all collective agreements. According to the Austrian Federation of Trade Unions (ÖGB), in a typical year, there are more than 400 collective agreements (without supplemental agreements at the firm level which do not cover wages). Source: *Jahresbericht des ÖGB* 2005.

classified TLI units into three "qualities": 86 units were assigned to quality 1 (the best category), 14 to quality 2 and 25 to quality 3. In our estimations we only used those 100 series with quality 1 or 2. See Table C-1 for the "quality" assigned to all units.

The more agreements are subsumed by an individual TLI series, the more likely it is that we cannot use it in our analysis because wage negotiations may take place in different months and we cannot identify the variables that we are interested in, i.e. the month of the wage change, the associated wage change and the contract length. A typical example is the index for blue-collar workers in the food industry which increases several times a year and was thus classified as "quality 3". The same is true for white-collar workers in the same industry.[30] Smaller units which we considered not to use include "blue-collar workers in other trades" (the name indicates that it is a composite of different branches), "blue-collar workers in the clothing industry" where there are several wage changes in almost every year and the associated months show no regular pattern. Finally, we also classified the index series of workers employed by professionals (such as doctors, lawyers and chemists) as "quality 3" because these are not included in the TLI 1976.

C.1.3 Panel structure of the data

If we use only index series with quality 1 or 2 we would have 2700 (100*27) wage change observations if there were exactly one wage change per year. However, the panel is not completely balanced. 27 annual wage change observations are available only for 49 units. In all other series at most 3 observations are missing. One reason for such gaps is that sometimes an agreement lasts longer than a year which could mean that no wage change is observable in a particular year (the most important example being the aforementioned three-year public-sector wage freeze in the 1990s which affected 19 series). We end up with 2621 wage change observations for 100 units in 27 years.

C.1.4 "Independent" TLI units

Although wage negotiations in many sectors are conducted separately for blue- and white-collar workers they are very often synchronous and the wage changes are very similar. The same holds for collective agreements of similar or related sectors. One could suspect wage negotiations in such cases are in effect strongly interwoven (as casual observations of

[30] On the other hand, the indices for the white-collar workers in wholesale or retail trade and the one for blue-collar workers in "supplementary construction trades" are usable because apparently the underlying collective agreements are synchronous.

current wage settlements also suggests). The same also holds for wages in the public sector.

As a robustness check, based on a subjective assessment of whether there is a high correlation *and* near-synchrony of wage changes (paired with an obvious relatedness between units), we grouped a number of series together and let each of these groups be represented by one of them (mostly the largest unit). For example, in construction four series were merged (construction industry and construction trade, both for blue-collar and white-collar workers). Note especially that four series of the metal and mining industries (again, for both blue-collar and white-collar workers) were grouped together. These constitute the group of wage-leading units (see below). Most importantly, 18 index series belonging to the public sector were also treated as one unit in this exercise. Altogether, these aggregations reduced the number units from 100 to 55. Our estimation results were hardly affected by this, as column (5) in Table 8 suggests.

C.2 Other data

C.2.1 Macro forecast data

We assume that expectations over inflation and real activity are shaped by forecast data of the Austrian Institute of Economic Research (WIFO). The WIFO is Austria's oldest economic research institute and its forecasts are said to be especially relevant for wage setting. This is plausible, because it is (partly) owned by the social partners and has close ties to them. We use series for real GDP growth, inflation (measured by the consumer price index) and the unemployment rate (based on registered employment and unemployment data).[31]

The WIFO projections have for a long time been published regularly at the end of each quarter. Typically, forecasts are made for the current as well as for the next year. However, until 1980, next-year forecasts were only provided with the September and December projections. Between 1981 and 1988, next-year forecasts were also published in June, and since 1989 each of the four annual projections (i.e. also the one in March) include both current- and next-year forecasts. In June 1997 there was no WIFO projection.

Our estimation equation (15) for each wage change is based on the most recent forecast available in each month of every year to construct a forecast *exactly one year ahead.* However, this would not be possible for many months in the period between 1980 to 1988

[31]We thank Josef Baumgartner for providing us with the forecast series.

because next-year forecasts are missing. So we had to make some assumptions how to fill in these "missing" forecast values. All forecasts for June 1997 were imputed as the simple averages of March and September 1997. To get the needed next-year forecasts for the years between 1980 and 1988 we experimented with two different approaches: In the first one we assumed that the forecast for the next year is equal to the forecast for the current year. In the second approach the assumption was that the next-year forecast of a variable is a weighted average of this year's forecast for that variable and that of the next year in September (where we have next-year forecasts throughout all observations). The weights are 0.75 for the current year and and 0.25 for next year for projections in March and 0.5 for the current year and 0.5 for the next year for projections in June, respectively. In the empirical exercises we sticked to the second approach but did not find substantial differences between both imputation approaches.

Table C-2: Imputations of WIFO macroeconomic forecasts

projection for:	current year				next year			
year / projection of:	March	June	September	December	March	June	September	December
1980	✓	✓	✓	✓	○	○	✓	✓
1981	✓	✓	✓	✓	○	✓	✓	✓
1982	✓	✓	✓	✓	○	✓	✓	✓
1983	✓	✓	✓	✓	○	✓	✓	✓
1984	✓	✓	✓	✓	○	✓	✓	✓
1985	✓	✓	✓	✓	○	✓	✓	✓
1986	✓	✓	✓	✓	○	✓	✓	✓
1987	✓	✓	✓	✓	○	✓	✓	✓
1988	✓	✓	✓	✓	○	✓	✓	✓
1989	✓	✓	✓	✓	✓	✓	✓	✓
1990	✓	✓	✓	✓	✓	✓	✓	✓
1991	✓	✓	✓	✓	✓	✓	✓	✓
1992	✓	✓	✓	✓	✓	✓	✓	✓
1993	✓	✓	✓	✓	✓	✓	✓	✓
1994	✓	✓	✓	✓	✓	✓	✓	✓
1995	✓	✓	✓	✓	✓	✓	✓	✓
1996	✓	✓	✓	✓	✓	✓	✓	✓
1997	✓	○	✓	✓	✓	○	✓	✓
1998	✓	✓	✓	✓	✓	✓	✓	✓
1999	✓	✓	✓	✓	✓	✓	✓	✓
2000	✓	✓	✓	✓	✓	✓	✓	✓
2001	✓	✓	✓	✓	✓	✓	✓	✓
2002	✓	✓	✓	✓	✓	✓	✓	✓
2003	✓	✓	✓	✓	✓	✓	✓	✓
2004	✓	✓	✓	✓	✓	✓	✓	✓
2005	✓	✓	✓	✓	✓	✓	✓	✓
2006	✓	✓	✓	✓	✓	✓	✓	✓

○ ... forecast not available (imputed)
✓ ... forecast available

C.2.2 Month-specific growth rates, forecasts and forecast errors

As different sectors reach agreements in different months of a particular year it is important to construct month-specific macroeconomic variables. As far as past values are concerned the change of a macro variable x in a particular month over the last year is computed by weighted annual averages: $\Delta x_{j,\tau} = \frac{j-1}{12}\Delta x_{\tau} + \frac{12-j+1}{12}\Delta x_{\tau-1}$ where j and τ are subscripts for month and year, respectively. For example, the inflation rate over the last 12 months in May 2003 is given by $\Delta p_{5,2003} = \frac{4}{12}\Delta p_{2003} + \frac{8}{12}\Delta p_{2002}$. Equivalently, we deal with GDP growth and changes of the unemployment rate.

As far as forecasts are concerned we have to be careful to represent as exactly as possible the expectations wage setters had at that time. We use a similar procedure as for the past values. In this case we define by $\Delta x^{fc}_{j,\tau}$ the expected change of x over the upcoming year starting in month j of year τ. Again, we use weighted yearly averages. This means that $\Delta x^{fc}_{j,\tau} = \frac{12-j+1}{12}\Delta x^{fc}_{\tau} + \frac{j-1}{12}\Delta x^{fc}_{\tau+1}$ where Δx^{fc}_{τ} and $\Delta x^{fc}_{\tau+1}$ are given by the WIFO projections for Δx for the years τ and $\tau + 1$, respectively. As there are four projections in each year (see above) we had to decide which of them is relevant in a particular month j. We proceeded as follows:

Table C-3: Allocation of WIFO projections to months of wage changes

Month of wage change	Relevant WIFO projection
1, 2, 3	December (last year)
4, 5, 6	March (current year)
7, 8, 9	June (current year)
10, 11, 12	September (current year)

In months 2 and 3 of a year τ there is a further difficulty, namely how to get a forecast for a full year ahead because the only macroeconomic projection available is the one from December in the previous year containing only forecasts for $\tau - 1$ and τ. However one would also need a forecast for $\tau + 1$ which is only available with the next (March) projection. In such a case we simply set the forecast for $\tau + 1$ equal to that for τ.

One could object it is implausible that for a wage settlement that becomes effective, say, in April the relevant projection is from the end of March. However, in most cases, not very much time passes between the point in time a settlement is reached and when it becomes effective (this time span is usually less than a month). It is also likely that wage setters are influenced by “rumours” about the outcome of the pending WIFO macroeconomic projection.

We use these constructed values $\Delta\tilde{x}^{fc}_{j,\tau}$ as proxies for the month-specific expectations of inflation ($E_{j,\tau}\Delta\tilde{p}_{j,\tau}$) and real activity ($E_{j,\tau}\Delta\tilde{y}_{j,\tau}$) in equation (15). Finally, after having constructed one-year-ahead forecasts and past-year-changes for the relevant macro variables in each month forecast errors are defined as $\Delta x^{err}_{j,\tau} = \Delta x_{j,\tau} - \Delta x^{fc}_{j,\tau-1}$.

C.3 Construction of reference norms

C.3.1 External reference norm

Call $D(i, j, \tau)$ the date when unit i sets its new wage in month j in year τ. The external reference norm is assumed to be given by the (weighted) average wage increase of all other (i.e. excluding i) units that have set new wages since $D(i, j', \tau')$, that is, since the last time the unit i has set its wage. Normally, j' is identical or close to j and, usually, $\tau' = \tau - 1$ but, as explained, the time distance could be up to three years.

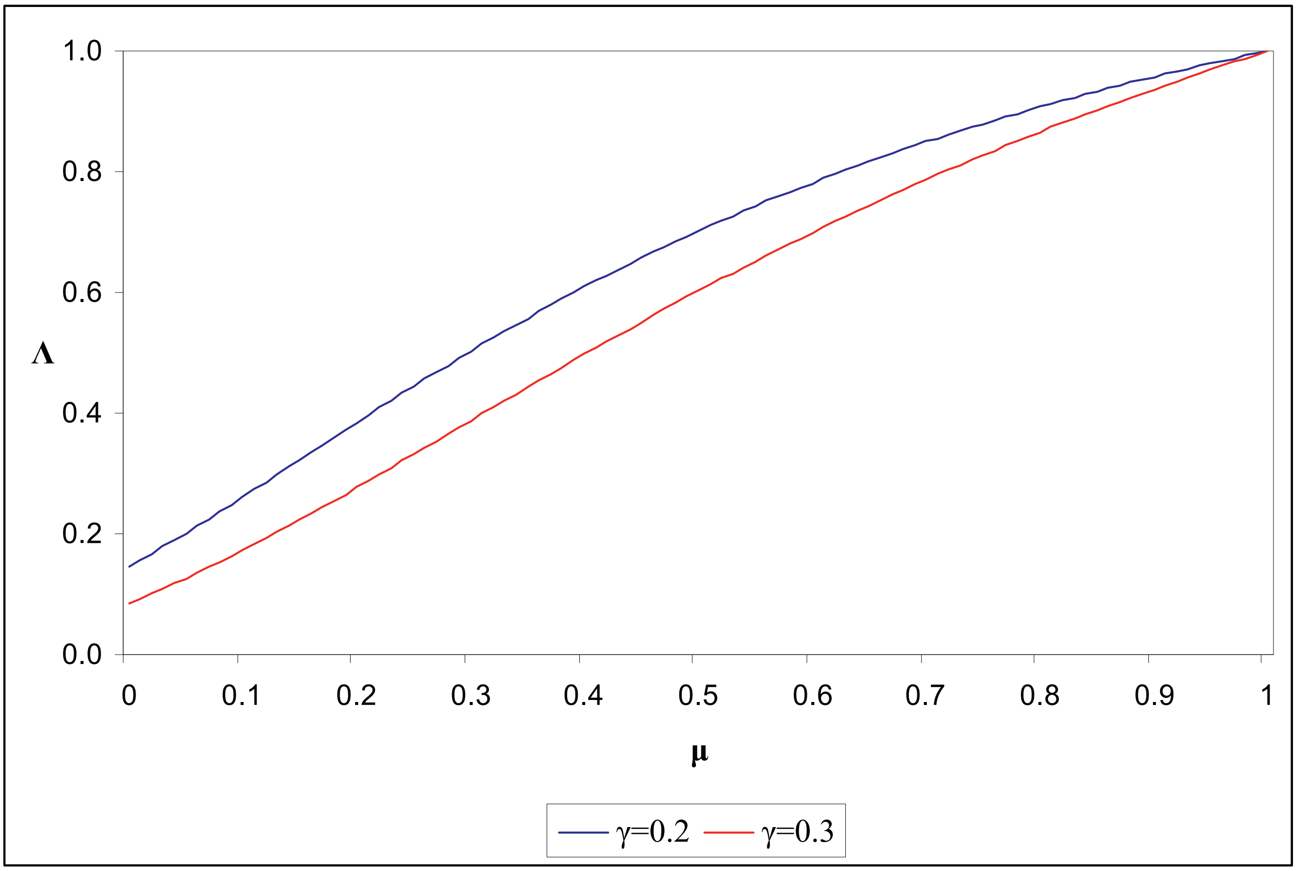
C.3.2 Wage leadership norm

This is simply the wage increase of the blue-collar workers[32] in the metal industry whereas for the wage-leading units themselves (blue-collar/white-collar workers in the metal and the mining industry, respectively) this norm is set to zero. If wage negotations for a TLI unit take place in November or December then we use the metal-sector wage change of the current year, in all other months (January to October) last-year wage changes.

[32]There are two series for the metal industry as the TLI also contains the one for white-collar workers. Wage increases are most often, but not always identical. Given the higher union density and power of the metal workers' union (which represents the blue-collar workers of that industry whereas white-collar workers are organized in another union) we chose the series of blue-collar workers as the wage leader.

Figures and Tables

Figure 1 – The Effect of the Importance of Reference Norms on Inflation Persistence Λ



Note: The figure reports the degree of annual inflation persistence Λ as the importance of reference norms increases from $\mu^A = \mu^B = \mu = 0$ to $\mu = 1$ for γ =0.2 and γ =0.3.

<u>Table 1</u> – The Effect of Different Reference Norms on Inflation Persistence Λ

	μ =0	μ =0.25	μ =0.5
		A. External Norms	
γ=0.2	0.146	0.444	0.702
γ=0.3	0.085	0.332	0.602
		B. Price Indexation Norms	
γ=0.2	0.146	0.314	0.539
γ=0.3	0.085	0.216	0.425
		C. Habit-Persistence Norms	
γ=0.2	0.146	0.721	0.916
γ=0.3	0.085	0.648	0.884

Note: The numbers report the degree of annual persistence $\Lambda = \lambda^A \lambda^B$ for three alternative reference norms (explained in the text) and various degrees of real rigidity (γ) and the importance of reference norms (μ). For the habit persistence norm the measure of persistence is the sum of the first two autoregressive terms.

<u>Table 2</u> – The Effect of Asymmetries in Reference Norms on Inflation Persistence Λ

	No Reference Norms ($\bar{\mu} = 0$)	***Symmetric Reference Norms*** ($\bar{\mu} = 0.5$)	***Asymmetric Reference Norms*** ($\bar{\mu} = 0.5$)
	$\mu^A = \mu^B = 0$	$\mu^A = \mu^B = 0.5$	$\mu^A = 0, \mu^B = 1$
γ=0.2	0.146	0.702	0.5
γ=0.3	0.085	0.602	0.368

Note: The numbers report the degree of annual persistence $\Lambda = \lambda^A \lambda^B$ when wage-setters have external reference norms as explained in the text.

Table 3 – Summary Statistics on Individual Collective Wage Agreements in Austria, 1980-2006

	Growth Rate of Wages		Quarter of New Wage Agreements					Length of New Agreements (in months)			
Year	Mean	Std. Dev.	Winter	Spring	Summer	Fall	No Contract	12	<12	13-24	>24
1980	0.0596	0.0207	44%	24%	13%	17%	2%	76%	3%	19%	2%
1981	0.0767	0.0209	45%	24%	12%	16%	3%	87%	1%	9%	3%
1982	0.0647	0.0114	47%	24%	11%	16%	2%	58%	4%	36%	2%
1983	0.0463	0.0098	43%	28%	12%	16%	1%	60%	26%	13%	1%
1984	0.0445	0.0095	43%	30%	9%	16%	2%	89%	1%	8%	2%
1985	0.0530	0.0059	46%	29%	8%	17%	0%	89%	1%	10%	0%
1986	0.0449	0.0111	47%	28%	8%	13%	4%	87%	2%	7%	4%
1987	0.0306	0.0052	48%	30%	8%	13%	1%	65%	2%	32%	1%
1988	0.0281	0.0096	24%	31%	32%	12%	1%	64%	26%	9%	1%
1989	0.0377	0.0146	49%	30%	6%	14%	1%	95%	0%	4%	1%
1990	0.0620	0.0160	48%	31%	7%	14%	0%	93%	3%	4%	0%
1991	0.0665	0.0095	49%	31%	7%	13%	0%	98%	1%	1%	0%
1992	0.0536	0.0089	49%	31%	7%	13%	0%	94%	0%	6%	0%
1993	0.0440	0.0166	49%	26%	7%	17%	1%	90%	6%	3%	1%
1994	0.0324	0.0078	51%	28%	7%	13%	1%	92%	2%	5%	1%
1995	0.0329	0.0081	49%	30%	6%	12%	3%	66%	4%	7%	23%
1996	0.0200	0.0127	33%	28%	4%	14%	21%	62%	2%	14%	22%
1997	0.0178	0.0122	26%	30%	6%	14%	24%	62%	5%	9%	24%
1998	0.0225	0.0066	47%	31%	7%	15%	0%	85%	6%	9%	0%
1999	0.0238	0.0067	47%	30%	4%	19%	0%	94%	3%	3%	0%
2000	0.0198	0.0071	48%	28%	4%	17%	3%	93%	2%	2%	3%
2001	0.0298	0.0356	50%	29%	4%	16%	1%	90%	2%	7%	1%
2002	0.0217	0.0082	49%	31%	4%	13%	3%	91%	2%	4%	3%
2003	0.0220	0.0064	50%	33%	4%	11%	2%	94%	0%	4%	2%
2004	0.0199	0.0055	52%	32%	5%	9%	2%	93%	1%	4%	2%
2005	0.0244	0.0051	51%	33%	6%	10%	0%	94%	3%	2%	1%
2006	0.0250	0.0045	53%	32%	5%	9%	1%	–	–	–	–
Total	0.0379	0.0212	46%	29%	9%	13%	3%	80%	4%	9%	7%

Note: The numbers in the table refer to the sample of 100 wage-setting units comprising 92% of the total labor force. The numbers are unweighted. The quarters are defined as follows: Winter (Jan., Feb., Mar.), Spring (Apr., May, Jun.), Summer (Jul., Aug., Sep.), Fall (Oct., Nov., Dec.). The length of new agreements refers to the year when they start.

Table 4a – Determinants of Collective Wage Agreements (External Reference Norms)

	Dependent Variable: growth rate of unit-specific wage rates ($\Delta w^{i}_{j,\tau}$)							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Estimation Method	FE	RE	RC	MG	FE	RE	RC	MG
External Norm	0.551***	0.548***	0.571***	0.551***	0.636***	0.635***	0.655***	0.650***
	(0.025)	(0.026)	(0.039)	(0.034)	(0.024)	(0.024)	(0.035)	(0.030)
Inflation (forecast)	0.595***	0.590***	0.550***	0.588***	0.578***	0.573***	0.545***	0.578***
	(0.034)	(0.034)	(0.041)	(0.035)	(0.032)	(0.031)	(0.042)	(0.037)
GDP growth (forecast)	0.328***	0.326***	0.327***	0.328***	–	–	–	–
	(0.036)	(0.035)	(0.046)	(0.042)				
Change in unemployment rate (forecast)	–	–	–	–	-1.226***	-1.240***	-1.150***	-1.212***
					(0.094)	(0.093)	(0.11)	(0.099)
Time (Decade) Dummies	YES	YES	YES	YES	YES	YES	YES	YES
Constant	-0.00801***	-0.00767***	-0.00741***	-0.0078***	-0.00144*	-0.00119	-0.00154*	-0.0021***
	(0.0010)	(0.0011)	(0.0012)	(0.001)	(0.00081)	(0.00086)	(0.00092)	(0.00068)
Observations	2621	2621	2621	2621	2621	2621	2621	2621
Number of groups	100	100	100	100	100	100	100	100
Test-statistic for Random Effects	–	83.04	–	–	–	78.38	–	–-
Probability value		0.000				0.000		
χ^2-statistic for H_0: Parameter Constancy	–	–	1664.14	–	–	–	1623.08	–
Probability value of H_0	–	–	0.000	–	–	–	0.000	–
Sum: $\hat{\beta}_{infl} + \hat{\beta}_{norm}$	1.146	1.138	1.121	1.139	1.214	1.208	1.2	1.228
F-statistic for H_0: $\hat{\beta}_{infl} + \hat{\beta}_{norm} = 1$	49.506	41.38	28.296	8.089	94.834	87.07	68.31	23.172
Probability value of H_0	0.000	0.000	0.000	0.005	0.000	0.000	0.000	0.000

Note: The tables contain the result of panel estimations of the determinants of unit-specific collective wage agreements $\Delta w^{i}_{j,\tau}$ in Austria, where i=1, 2, …, 100, j=1, 2, …, 12 and t=1980, 1981, …, 2006. The estimation uses the assumption of external reference norms and GDP growth (columns (1) to (4)) or the change in the unemployment rate (columns (5) to (8)) as a measure of real activity. The time dummies are defined as decade dummies (i.e. for 1980-1989, 1990-1999, etc.). For each model we report the results of a fixed effects (FE), a random effects (RE), a random coefficients (RC) and a mean group (MG) estimation. In columns (2) and (6) we report the Hansen-Sargan statistic for overidentifying restrictions in order to test for the appropriateness of the RE specification. For the RC estimation we report the χ^2 statistics for parameter constancy proposed by Swamy (1970). For all specifications we also give the sum of the coefficients on the norm and the inflation forecasts and also the F-statistic and the p-value of testing whether this sum is equal to 1. Standard errors are in parentheses. ***, ** and * denote statistical significance at the 1, 5 and 10 % level, respectively.

Table 4b – Determinants of Collective Wage Agreements (Wage Leadership Reference Norms)

	Dependent Variable: growth rate of unit-specific wage rates ($\Delta w^{i}_{j,\tau}$)							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Estimation Method	FE	RE	RC	MG	FE	RE	RC	MG
Leadership Norm	0.566***	0.346***	0.578***	0.599***	0.561***	0.348***	0.579***	0.594***
	(0.018)	(0.027)	(0.025)	(0.021)	(0.017)	(0.026)	(0.026)	(0.022)
Inflation (forecast)	0.496***	0.715***	0.439***	0.436***	0.519***	0.739***	0.454***	0.466***
	(0.026)	(0.037)	(0.031)	(0.025)	(0.027)	(0.037)	(0.034)	(0.028)
GDP growth (forecast)	0.0919***	0.187***	0.0751*	0.0568	–	–	–	–
	(0.032)	(0.036)	(0.042)	(0.038)				
Change in unemployment rate (forecast)	–	–	–	–	-0.365***	-0.591***	-0.219**	-0.260***
					(0.086)	(0.094)	(0.10)	(0.088)
Time (Decade) Dummies	YES	YES	YES	YES	YES	YES	YES	YES
Constant	-0.00270***	-0.00134	-0.00223**	-0.0024***	-0.000428	0.00306***	-0.000749	-0.0024***
	(0.00093)	(0.00100)	(0.0011)	(0.00086)	(0.00071)	(0.00086)	(0.00093)	(0.00069)
Observations	2621	2621	2621	2621	2621	2621	2621	2621
Number of groups	100	100	100	100	100	100	100	100
Test-statistic for Random Effects	–	485.36	–	–	–	461.91	–	–
Probability value		0.000				0.000		
Test statistic for H_0: Parameter Constancy	–	–	3317.48	–	–	–	3068.64	–
Probability value of H_0	–	–	0.000	–	–	–	0.000	–
Sum: $\hat{\beta}_{infl} + \hat{\beta}_{norm}$	1.062	1.062	1.017	1.035	1.081	1.088	1.033	1.06
F-statistic for H_0: $\hat{\beta}_{infl} + \hat{\beta}_{norm} = 1$	10.466	8.158	0.595	1.188	15.183	15.259	1.866	2.853
Probability value of H_0	0.001	0.004	0.44	0.276	0.000	0.000	0.172	0.091

Note: The tables contain the result of panel estimations of the determinants of unit-specific collective wage agreements $\Delta w^{i}_{j,\tau}$ in Austria, where i=1, 2, …, 100, j=1, 2, …, 12 and t=1980, 1981, …, 2006. The estimation uses the assumption of leadership reference norms and GDP growth (columns (1) to (4)) or the change in the unemployment rate (columns (5) to (8)) as a measure of real activity. The time dummies are defined as decade dummies (i.e. for 1980-1989, 1990-1999, etc.). For each model we report the results of a fixed effects (FE), a random effects (RE), a random coefficients (RC) and a mean group (MG) estimation. In columns (2) and (6) we report the Hansen-Sargan statistic for overidentifying restrictions in order to test for the appropriateness of the RE specification. For the RC estimation we report the χ^2 statistics for parameter constancy proposed by Swamy (1970). For all specifications we also give the sum of the coefficients on the norm and the inflation forecasts and also the F-statistic and the p-value of testing whether this sum is equal to 1. Standard errors are in parentheses. ***, ** and * denote statistical significance at the 1, 5 and 10 % level, respectively.

Table 5 – Comparison of Different Reference Norms with J-Tests

	t-statistic on the fitted values from a regression including			
	Leadership Norm	**External Norm**	**Habit Norm**	**Price Indexation Norm**
In regression including:				
Leadership Norm	–	3.592 (0.000)	2.804 (0.000)	0.332 (0.74)
External Norm	6.996* (0.000)	–	2.07 (0.038)	0.648 (0.517)
Habit Norm	20.481* (0.000)	15.357* (0.000)	–	3.123 (0.002)
Price Indexation Norm	20.955* (0.000)	16.09* (0.000)	10.946* (0.000)	–

Note: The values in the table are based on an approach where the fitted value of a benchmark regression with the respective reference norm in the column is added to a regression that includes the reference norm in the respective row. The numbers reported are the t-statistic of these fitted values while p-values are shown in parentheses. A "*" indicates that the t-statistic in one pair of comparisons is higher than in the case where the role of the two reference norms is reversed.

Table 6 – Nested Comparison of Different Reference Norms

	Dependent Variable: growth rate of unit-specific wage rates (Δw_t^i)					
	(1)	(2)	(3)	(4)	(5)	(6)
Leadership Norm	0.412*** (0.059)	0.563*** (0.027)	0.577*** (0.028)	–	–	–
External Norm	0.279*** (0.078)	–	–	0.718*** (0.047)	0.661*** (0.041)	–
Habit Norm	–	0.065* (0.023)	–	-0.065** (0.032)	–	0.307*** (0.028)
Price Indexation Norm	–	–	-0.016 (0.048)	–	-0.037 (0.056)	0.178*** (0.057)
Inflation (forecast)	YES	YES	YES	YES	YES	YES
Change in unemployment rate (forecast)	YES	YES	YES	YES	YES	YES
Time (Decade) Dummies	YES	YES	YES	YES	YES	YES
Constant	YES	YES	YES	YES	YES	YES

Note: The table contains the results if two reference norms are included in pairs into the benchmark estimation (column (7) of Table 4b).

Table 7 – Rejection Rate for the Individual Coefficients

	Significantly different from 1 (5% level)	Significantly different from 1 (1% level)
Leadership Norm	15%	6%
External Norm	69%	51%
Habit Norm	89%	60%
Price Indexation Norm	27%	11%
No Reference Norm	23%	9%

Note: The table contains the percentage of the 100 wage-setting units for which the condition $\hat{\beta}^i_{\inf} + \hat{\beta}^i_{norm} = 1$ on their individual coefficients is violated. The results are based on the RC estimation with the change in the unemployment rate as the measure of real activity.

Table 8 – Robustness Tests for the Benchmark Estimation (leadership reference norms)

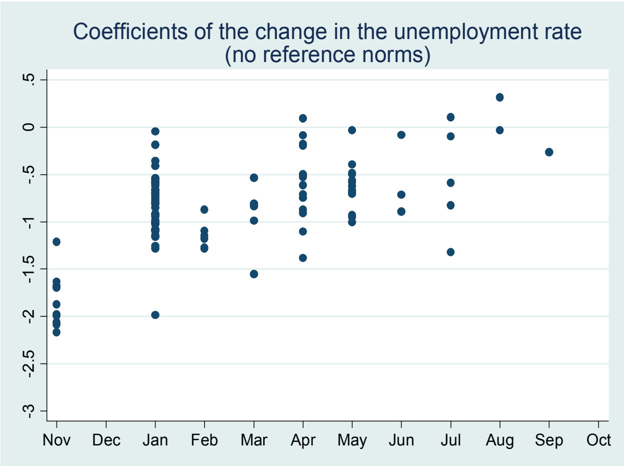
	Dependent Variable: growth rate of unit-specific wage rates (Δw_t^i)									
Estimation Method	RC	RC	RC	RC	RC	RC	RC	RC	RC	RC
	Benchmark (1)	<1993 (2)	>=1993 (3)	Only priv. sector (4)	Only indep. units. (5)	Forecast Errors (6)	Contract Length (7)	Past Inflation (8)	Level of Unr (9)	Level and Lag of unr (10)
Leadership Norm	0.579***	0.517***	0.436***	0.608***	0.625***	0.570***	0.583***	0.577***	0.617***	0.625***
	(0.026)	(0.024)	(0.099)	(0.030)	(0.035)	(0.042)	(0.026)	(0.028)	(0.031)	(0.031)
Inflation (forecast)	0.454***	0.458***	0.531***	0.437***	0.416***	0.460***	0.449***	0.479***	0.402***	0.482***
	(0.034)	(0.039)	(0.068)	(0.039)	(0.046)	(0.060)	(0.034)	(0.054)	(0.041)	(0.045)
Change in unemployment rate (forecast)	-0.219**	-0.326***	-0.105	-0.230*	-0.152	-0.361**	-0.192**	-0.225**		
	(0.10)	(0.13)	(0.17)	(0.12)	(0.15)	(0.16)	(0.084)	(0.11)		
Forecast error (inflation)						0.0171				
						(0.061)				
Forecast error (Δunr)						0.146				
						(0.17)				
Length of Contract (months)							0.001**			
							(0.00056)			
Lagged Inflation								-0.016		
								(0.048)		
Level of unemployment rate (forecast)									0.005	-0.207**
									(0.078)	(0.10)
Level of unemployment rate (current)										0.317***
										(0.089)
Constant, Time Dummies	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Observations	2621	1283	1338	2090	1446	2621	2621	2621	2621	2621
Number of groups	100	100	100	79	55	100	100	100	100	100
Sum: $\hat{\beta}_{infl} + \hat{\beta}_{norm}$	1.033	0.976	0.968	1.045	1.041	1.03	1.032	1.056	1.019	1.107
Probability value of $H_0 \left(\hat{\beta}_{infl} + \hat{\beta}_{norm} = 1\right)$	0.172	0.469	0.66	0.107	0.277	0.334	0.114	0.318	0.707	0.045

Note: The columns contain various robustness tests to the benchmark estimation with leadership reference norms in column (7) of Table 4b, here repeated in column (1). All estimations include (decade) time dummies, a constant term and they are based on RC models. Standard errors are in parentheses. ***, ** and * denote statistical significance at the 1, 5 and 10 % level, respectively.

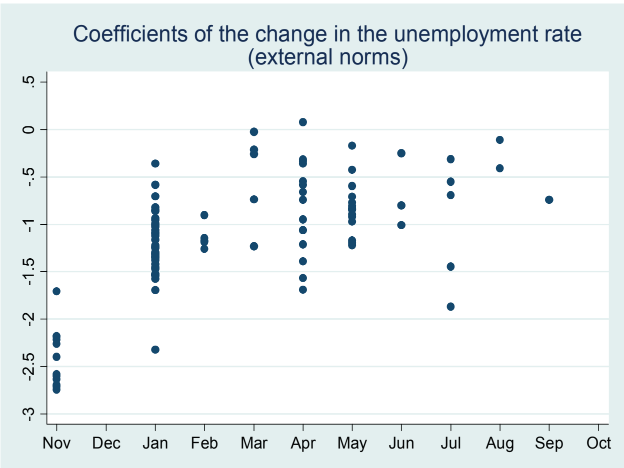
Figure 2 – The Reaction of the Individual Wage-Setting Units

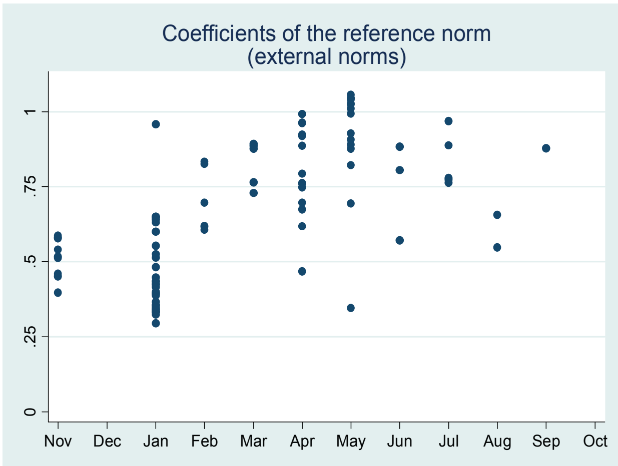
Coefficients of the expected change in the unemployment rate $\left(\hat{\beta}_2^i\right)$	Coefficients of the reference norm $\left(\hat{\beta}_3^i\right)$

Panel A: No Reference Norms

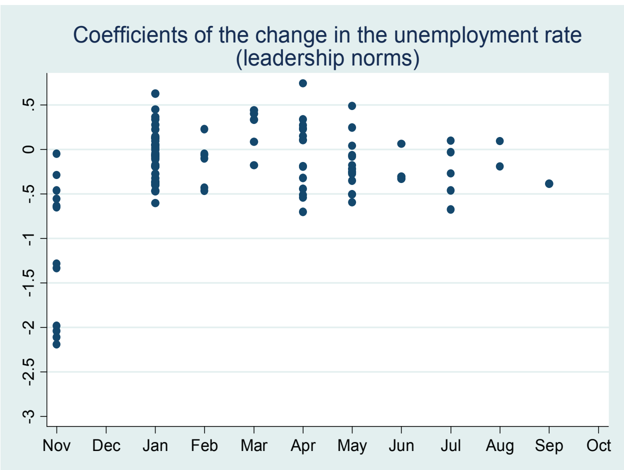


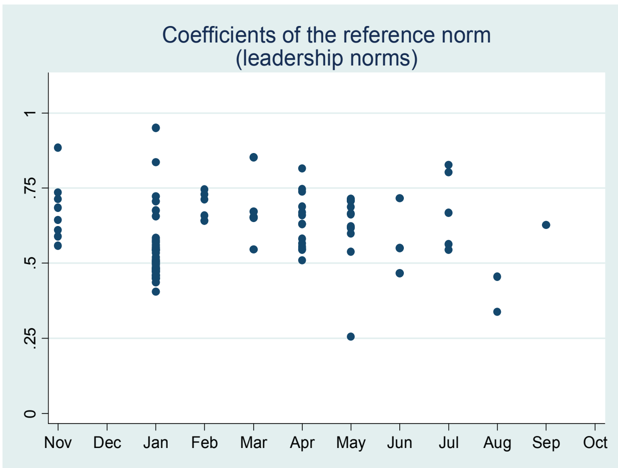
Panel B: External Reference Norms





Panel C: Leadership Reference Norms





Note: The graphs report the coefficients for the expected change in the unemployment rate and the reference norm for the 100 individual wage-setting units when the benchmark equation is estimated with a random coefficients model. The individual coefficients are ordered according to the typical (median) month in which each wage-setting unit has concluded its wage agreements.

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