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Have Inflation Targeting and EU labour Immigration Changed the System of Wage Formation in Norway

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Have inflation targeting and EU labour immigration changed the system of wage formation in Norway?*

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Abstract

Collective agreements have played a central role in the system of wage formation in Norway for more than fifty years. Although the degree of coordination achieved has been variable, pattern wage bargaining has been a mainstay of the system. We investigate the degree of invariance in wage formation in Norway with respect to two recent structural changes: the transition towards inflation targeting in monetary policy and an unprecedented surge in labour supply due to higher immigration rates. We report empirical results that support the view that a semi-permanent high immigration may affect wages negatively in a significant way. However we do not find evidence that the stability of the arbitration system, and in particular the wage-bargaining pattern, has been changed by labour immigration or by inflation targeting monetary policy. An explanation of why we do not find evidence of structural changing effects of the transition of monetary policy, can be found in the fact that the wage arbitration system itself has syncronized the inflation expectations of the social partners. In that analysis, inflation targeting became a new layer of nominal stabilization, on top of the existing one.

Keywords: *Inflation modelling, pattern wage bargaining, inflation targeting, dynamic econometrics, cointegration, small open economy*

JEL classification: *C52, E24, E31, E37, J31.*

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1 Introduction

In Norway, the post-war system for wage formation was originally designed to balance the need for relative inflation control with the attainment of full employment and a balanced current account, cf. Aukrust (1977). The system has survived both the international stagflation of the 1970s and the Norwegian housing and banking crisis in 1989-90. It was based on an agreement between the government and the wage setters. The government committed to a fixed exchange rate policy, while the wage setters committed to collective wage bargaining. The system of wage fixing in Norway unfolds along both the horizontal and vertical dimension, implying that wage setting is based both on bargaining between evenly matched labour unions and firms' organizations, and bargaining at a high level. Later, similar systems were adopted in, for example, Sweden, see Edgren et al. (1969) and the Netherlands. A presentation of the Norwegian model from an econometric perspective is found in Bårdsen et al. (2005, Ch. 3).

The system for wage formation created a link between the wage norm for the overall economy, which was determined by the sector exposed to international competition, and profitability and productivity growth in the same sector. For several decades the system has linked wage increases in the other sectors of the economy (known as wage-followers) to wage formation in the exposed sector (the wage-leader). Down the years, it has been accepted, even supported, by governments of different party colours. Probably a recognition of the idea that this version of pattern bargaining represents an operational way of creating a certain degree of coordination in wage setting.

One threat to this system is the secular decline in the traditional manufacturing industries, which in Norway has been accelerated by the extraction of oil and gas and the use of the resulting 'oil-money'. Nevertheless, several commentators have noted the stability of the Norwegian norm-based system. For example by Visser (2013, p. 59-60), and also by OECD economists:

> ... the so called traditional ('non-oil') sector has been diminishing in importance. Despite this, it is important in wage setting. Rather than wages being determined by the relative bargaining strength of different sectors, the general wage level is set by the social partners first considering the wage increases that the traditional sector can "afford". OECD (2012, p. 15)

Nevertheless, the functioning of the wage formation system has been challenged by two important events in the last decade: the transition towards inflation targeting formally introduced in 2001 and the EU enlargement in 2004, which led to a large increase in the labour supply.

The change in monetary policy, from exchange rate stabilization to inflation targeting, can potentially change wage formation and the degree of coordination among wage setters, see Cuikerman and Lippi (1999), Soskice and Iversen (2000) and Holden (2005). In Norway, the introduction of inflation targeting in 2001 could potentially have a direct impact on wage formation, to the extent that it came into conflict with the 'existing order' based on manufacturing sector wage leadership. This potential for conflict was brought into focus by the central bank's *Inflation Report* from 2002. That report stated that the new monetary

policy required a reversal of the causality flow in wage formation.[1] The bank's worry was that the wage leader model could pull inflation in one direction (up for example, because productivity growth is higher there than in the service sectors), at the same time as monetary policy analysis suggested a reduction of aggregate demand relative to the bank's estimate of the Norwegian economy's potential output. This issue can also be related to theoretical research, showing that the pattern of wage bargaining under inflation targeting may be of marginal importance, e.g. Calmfors and Seim (2013). Hence, the hypothesis emerged that new monetary policy would force a change of roles in the pattern wage bargaining, where the wage-followers may become wage-leaders and vice versa.

With regards to immigration, research based on microeconomic data has found evidence that immigration has reduced wage earnings in certain groups in the labour market, see Bratsberg and Raaum (2012). The hypothesis is that, unless the immigrants emulate incumbent workers as regards trade union membership, increased immigration may reduce trade unions' bargaining power. The result might be that wage growth in sectors with high immigration rates deviates from the general wage growth.

We investigate empirically the degree of invariance in wage formation with respect to these structural changes. First, we investigate whether the value of labour productivity remains the main determinant of the wage trend in the exposed sector. Second, we test whether the wage norm created by the exposed sector still has a defining role in Norwegian wage formation. Third, we investigate whether inflation expectations have become more important after the introduction of inflation targeting. Finally, we include a labour immigration variable in our information set and identify its role in our structural model of wage formation.

We use quarterly data from 1980 to 2011. This means that about 1/3 of the sample comes from the era with a monetary policy regime that targeted inflation, and that about 1/4 of the sample comes from the recent period of increasing and high labour immigration. Hence, if either the monetary policy regime shift or the new immigration flow has affected wage formation or the social order with any force, the evidence should be in this data set.

The paper is organised as follows. In Section 2, we outline the theoretical framework of coordination in Norway and define the twin concepts of wage-leader and wage-follower sectors in wage formation. Based on the theoretical and statistical properties of the time series in Section 3, we formulate an empirical model, test the wage-leader/-follower properties, and carry out analysis of the model in Section 4. In Section 5, we conclude that the econometric results suggest that the Norwegian wage model has preserved its main properties in spite of the recent challenges to the system.

[1] The following quotation from Inflation Report number 3/2002, page 28 summarizes the hypothesis: *With an inflation target for monetary policy and a floating exchange rate, it is the inflation target, not wage growth abroad, that determines the level of growth which is consistent with stable profitability in the business sector over time. Inflation in Norway will over time be determined by the inflation target that the Government has set for monetary policy and not inflation abroad. Exchange rate developments are determined by inflation differentials between Norway and other countries.*

2 Wage formation

The twin concepts of wage-leader and wage-followers are important in the Norwegian system of wage formation. Wage-leader refers to the sector (or sectors) where the bargaining outcome defines the wage norm that forms the basis for bargaining in the other sectors of the economy (wage-followers). Throughout the whole post-war period, the settlements in the manufacturing sector have in practice defined the wage norm.[2]

We also define the manufacturing industry as the wage-leading sector but we test whether that role has been preserved so far in the new millennium. For the wage-followers, we have chosen to distinguish between the private service sector (including construction, mainly because of the importance of engineering consultancy in this sector) and the public sector. But also because negotiations in the public sector start after the wage settlements in private sector. Hence, in the following, we number the sectors 1 (wage-leader), 2 (wage-follower, private) and 3 (wage-follower, public).

2.1 The Wage-leader

The formulation of a theory of a sustainable wage norm requires an assessment of not only self-interest among workers and firms, but also of compromise. As pointed out by Usher (2012), 'compromise is then not just another way of talking about self-interest, and social, political and institutional forces are not merely cover-ups for imprecisely modelled individuals rational actions, they are among the fundamental determinants of decisions. In this view, even a full analysis of rational behaviour leads to an indeterminacy of wages, and other considerations had to be introduced to resolve it.

The recognition among economists that there is an indeterminacy in the economic theory of wages goes back to the 1950s, see Forder (2014, Ch. 1.4) who cites Samuelson (1951, p. 312) and Hicks (1955, p. 390) and other leading theorists. The economic theory of supply and demand could set some limits to what wages can be set, but within those limits closer determination requires that other relationships are introduced. The indeterminacy of wages from theory also characterizes the now standard Diamond-Mortensen-Pissarides (DMP) search and matching model. In the DMP model, the wage is usually determined in a Nash bargaining game. But is the wage logically equal to the Nash solution given the assumptions of the DMP model? As Hall (2005) pointed out, any wage in the bargaining set is in principle consistent with private efficiency on the part of both the firm and the worker. In that sense, the equilibrium wage rate is only set-identified. He then went on to analyze a solution where the real wage is fixed, which however is only one possibility of what in the DMP-literature is referred to as wage 'stickiness'. [3]

[2] Because of its importance, the system of wage-formation is a recurrent theme in public (and at times academic) debate. For example, the "traditional" definition of wage-leader has been challenged several times over the last decades. Those wanting to "change the content" of the wage-leading sector has emphasized both reduced legitimacy of the norm, as the importance of the sector for the total wage bill has become smaller, and reduced relevance (since Norwegian manufacturing may have become integrated in the super profitable oil and gas extraction). Among the arguments that are used in favour of no-change is that a manufacturing wage-norm aids coordination.

[3] Following Hall (2005), several papers have incorporated rigid wage setting in search models. For instance, Gertler and Trigari (2009) present a DMP model where the frequency of wage bargaining is constrained by

At the same time as we find it challenging to determine wages theoretically, we also observe that actual wage bargains are struck year after year, and that they are rationalized by considerations of profits, actual and required (to attract investments), cost of living and relative wages (fairness). These observed regularities give reason to believe that wage formation can be subject to econometric treatment. This is also what have motivated much of the econometric literature recently surveyed by Forder (2014). A development that goes back to the first half of the 1960s in Norway (although it was documented in English only later, by Aukrust (1977)) was that the wage norm became defined as a long-run trend, and also that short-run deviations had to be tolerated as 'part of the system'. But how should the trend be defined? Should it be a combination of the many things that might legitimate a secular trend trend? Or should the number of norm determined factors be limited? In Norway, a view that won support on both sides for the arbitration system, was that it was important to 'keep it simple' and relate the wage norm to the value of labour productivity in the exposed sector of the economy.

This idea of a wage norm can be treated econometrically using cointegration methods and equilibrium correction modelling. One of the main implications that we follow up is that, measured on a logarithmic scale, the wage level in the exposed sector, should be cointegrated with the log of the level of product prices and average labour productivity. Long-run price homogeneity together with a unit long-run elasticity on productivity lead to the implication of a stationary wage-share which may serve as an operationalization of an equilibrium wage-share. In the wider interpretation, the equilibrium income distribution needs not be completely constant but can depend on intermittent shifts, or trends, in bargaining power (possibly proxied by the unemployment rate) and the support to the idea about necessary compromise and coordination.

We base our econometric model on the assumption that both product price (q_1) and average labour productivity (z_1) are random-walk processes with drift. It follows that the time series for w_{1t} (the logarithm of the hourly wage in the wage-leader sector) contains both a deterministic and a stochastic (random walk) trend.[4]

Formally, the value of average labour productivity can be split into two components: q_{1t} (the log of the price in Norwegian currency) and z_{1t} (the log of average productivity in fixed prices), and these exogenous processes can be expressed as two random-walks with drift:

$$q_{1t} = q_{10} + q_{1t-1} + v_{q_1 t} \tag{1}$$

$$z_{1t} = z_{10} + z_{1t-1} + v_{a_1 t} \tag{2}$$

(1) and (2) imply that q_1 and z_1 are integrated of order 1, denoted $I(1)$. Positive deterministic trends in prices and productivity require $q_{10} > 0$ and $z_{10} > 0$.

Note that we could write equation (1) in terms of a price that is denoted in foreign currency, q^f_{1t}, and the exchange rate xr_t (log of kroner per currency unit) and specify time

Calvo (1983)) style lottery, leading to sticky wages. Blanchard and Galí (2010) combine a reduced form of search model with real wage rigidity with a New Keynesian model to study how this impacts monetary policy. Krogh (n.d.) generalizes the Hall-approach to a small open economy model where there is a non-trivial distinction between the consumer real wage and the producer real wage.

[4]In this section, we abstract from payroll tax rates in all three sectors. A representative payroll tax is included in the empirical model.

series models for q^f_{1t} and xr_t. To be consistent with $q_{10} > 0$ in equation (1), one or both of these time series models must contain a positive drift parameter. We assume that the price level variable q^f_{1t} has a positive drift, and for simplicity, we also assume that it is a constant parameter. The specification of the process for the nominal exchange rate is more complicated. For the period with a fixed exchange rate regime, it can possibly be specified without drift, but with intermittent structural breaks in order to represent devaluations. After the switch to the present floating exchange rate regime, a random walk, possibly with drift, is probably reasonable for the nominal exchange rate.

Hence, also across exchange rate regimes, equation (1) with $q_{10} > 0$ is a reasonable first approximation of the evolution of the price level in the wage-leading sector. As noted, there may be structural breaks in the drift parameter, and the foreign exchange rate regime-shift may also mean v_{q_1t} is not white-noise. The main point, however, is that the unit root property is robust across regimes and, in particular, that there are no explosive roots that logically must be included in the model of wage-price formation during the period with a floating exchange rate.

As noted, we assume that q_{1t} and z_{1t} jointly determine the deterministic trend in w_{1t}. In the simplest case, w_{1t} is cointegrated with the sum $q_{1t} + z_{1t}$, meaning that the logarithm of the wage share is I(0) with a constant mean s_1. This relationship can be written as:

$$w_{1t} - q_{1t} - z_{1t} = s_1 + e_{1t} \tag{3}$$

$$e_{1t} \sim I(0), \tag{4}$$

The sum of the productivity trend and the foreign price trend plays an important role in the framework, since it traces out a central tendency or a long-run sustainable scope for wage growth. This wage norm seems to correspond well to the concept of a wage corridor (Aukrust (1977)) for wage determination in the industries that are most exposed to foreign competition.

Cointegration implies equilibrium correction. Therefore, with reference to equations (1) and (2), e_{1t-1} should have significant predictive power for wage growth, Δw_{1t} ($\equiv w_{1t} - w_{1t-1}$):

$$\begin{array}{lll} a) & (w_1 - q_1 - z_1)_{t-1} & \rightarrow \Delta w_{1t} \\ b) & (w_2 - w_1)_{t-1} & \nrightarrow \Delta w_{1t} \\ c) & (w_3 - w_1)_{t-1} & \nrightarrow \Delta w_{1t} \end{array} \tag{5}$$

The economic interpretation is that collective wage bargaining, through its focus on the distribution of value added between labour and capital, implies that an equilibrium wage share (the parameter s_1) is maintained over time, cf. Forslund et al. (2008). (5b) and (5c) capture the idea that if the manufacturing sector is wage leading, it is implied that Δw_{1t} cannot be influenced by the lagged relative wage to the two other sectors.

As mentioned above, the equilibrium wage share is likely to depend on several underlying factors, e.g. related to production technology, product market conditions and bargaining power. It is only when these factors are constant that s_1 is likely to be a stable parameter. More generally, we can therefore add a time subscript to s_1 in equation (3) and write s_{1t} as a function of the variables that we include as its determinants in the econometric model, namely the log of the unemployment rate, u_t, and the immigration rate, IM_t:

$$s_{1t} = s_{10} + \beta_{11} u_t + \beta_{12} IM_t.$$

Hence, the extended long-run wage equation for sector 1 that we use in the following becomes

$$w_{1t} - q_{1t} - z_{1t} = s_{10} + \beta_{11} u_t + \beta_{12} IM_t + e_{1t} \tag{6}$$

If we abstract from IM, (7) is a standard linearised wage bargaining model, with full weight on producer prices and no weight on consumer prices, see e.g. Nickell and Andrews (1983) and Hoel and Nymoen (1988). The signs of the parameters, β_{11} and β_{12}, are expected to be negative (or zero). A higher unemployment rate can reduce the bargaining power of the unions. A marked change in the immigration rate may also affect union bargaining power and coordination negatively, unless a large proportion of the immigrants choose to become union members.

2.2 Wage-followers

If the manufacturing wage level represents the wage norm in the pattern of wage bargaining, and if $w_{1t} \sim I(1)$, w_{2t} must also be $I(1)$ and cointegrated with w_{1t}:

$$w_{2t} - w_{1t} = s_2 + e_{2t}\ , \tag{7}$$
$$e_{2t} \sim I(0), \tag{8}$$

where s_2 is the equilibrium relative wage. Note that equation (7) and (8) can be maintained by adjustments to wages in both sector 1 and sector 2. In order to define sector 2 as a wage-follower with sector 1 as a wage-leader we require:

$$\begin{array}{llcl} a) & \Delta w_{1t} & \overset{\longrightarrow}{\not\leftarrow} & \Delta w_{2t} \\ b) & (w_{2t-1} - w_{1t-1}) & \longrightarrow & \Delta w_{2t} \end{array} \tag{9}$$

Requirement (9a) implies that the contemporaneous relationship between the wage growth rates in the two sectors is recursive. We can implement and test (9a) as a restriction on a simultaneous equations model of wage setting. (9b) requires that the stationarity of the relative wage is due to equilibrium correction in sector 2.

In the Norwegian system of wage fixing, bargaining in the public sector starts after negotiations in the private sector, and it is therefore natural to assume that w_{3t} equilibrium corrects with respect to w_{2t}, i.e.:

$$w_{3t} - w_{2t} = s_3 + e_{3t} \tag{10}$$
$$e_{3t} \sim I(0). \tag{11}$$

The required recursive structure is the same as above, but between sector 2 and sector 3:

$$\begin{array}{llcl} a) & \Delta w_{2t} & \overset{\longrightarrow}{\not\leftarrow} & \Delta w_{3t} \\ b) & (w_{3t-1} - w_{2t-1}) & \longrightarrow & \Delta w_{3t} \end{array} \tag{12}$$

Clearly, for the government sector to be follower of sector 2, Δw_2 in (9) should not respond to $(w_3 - w_2)_{t-1}$. In both sectors, we include the possibility of potential wage shifting in the empirical model. Hence, s_2 in equation (7) and s_3 in equation (10) can include effects from the unemployment rate and from the immigration flow in the same way as in equation (6).

2.3 Wage-price inflation

The wage-price spiral is well known from the literature, cf. Blanchard (1987), Meade (1982) and Layard et al. (1991). In the wake of the financial and jobs crisis, the dynamics of wage and price setting have come to be regarded as central to rebalancing the euro area after the crisis, cf. ECB (2012) and OECD (2014).

Wage increases to compensate for expected increases in the cost of living are regularly demanded in the collective wage bargaining context, also in the wage-leading manufacturing sector. In Norway, a government-supported body, the Technical Calculation Committee for Wage Settlements (TCC), has the anchoring of CPI inflation forecasts as one of its tasks. The committee consists of representatives of both employer and employee organisations, and the consolidation of expectations takes place before the annual rounds of wage fixing starts.

Since CPI inflation, in turn, depends on growth in wage costs (in addition to the price of imported consumer goods), it is clear that a realistic system of equations for Δw_{1t}, Δw_{2t} and Δw_{3t} will not be completely recursive: in the short run, wage adjustments in the manufacturing sector also depend on CPI inflation expectations.

Nevertheless, as long as the wage norm that defines e_{1t} does not include the level of CPI, the inclusion of the rate of inflation in the system of wage formation represents a realistic modification of the otherwise recursive structure.

2.4 Potential reversal of bargaining pattern

As noted in the introduction, it is possible that inflation targeting caused an inversion of the causal ordering in wage formation. Specifically, that the relative wage $w_1 - w_2$ becomes a predictor of Δw_1, rather than of Δw_2. According to this hypothesis, neither (5b) nor (9b) will hold empirically on the latter sample period when monetary policy has changed. The hypotheses of reversed causality can be tested within the cointegrated VAR .The VAR is estimated in section 4.

3 Temporal properties of the time series variables

In this section, we first examine the assumption that wages, prices and productivity are non-stationary. Second, we examine whether the residuals of the equilibrium correction relationships are stationary. Stationary residuals between integrated variables imply a cointegrating relationship between the variables in the long-term solution.

Time series for wages, productivity and producer price developments are taken from the National Accounts. The unemployment rate is from the labour force survey, while consumer prices, immigration and population figures are collected from the official pages of Statistics Norway. Variables measured in natural logarithms are denoted by lower case Latin letters. Appendix A contains a detailed variable description.

3.1 The manufacturing sector

Wages, product prices and productivity in the manufacturing sector have increased over time, see Figure 1. The statistical properties of the variables are tested using two unit

root tests, the Augmented Dickey Fuller test (ADF) and a test by Kwiatowski et al. (1992) referred to as KPSS henceforth. The null of the ADF test is a unit root, and a rejection of the null implies that the time series is stationary. The null of the KPSS, on the other hand, is stationarity. Hence a rejection implies that the time series is integrated. Both tests are used to shed light on the time series properties.

The KPSS test rejects that w_1 is stationary, see Table 1. The ADF test statistic (including a constant, trend and seasonal dummies) is just above the critical value, which suggests the opposite result. The KPSS test result is supported by the testing of stationarity of the first difference of the same variable. The ADF test statistic of w_1 in difference (including a constant and seasonal dummies) is clearly above the critical value. Results from both ADF and KPSS tests indicate that the time series for product prices, q_1 is integrated. The ADF test statistic for productivity, z_1, does not reject I(1), while the KPSS test indicates that stationarity cannot be rejected. While the opposite is true for consumer prices, p_t, the ADF test rejects I(1), while the KPSS test strongly rejects stationarity. However, it is logically inconsistent to regard CPI inflation as I(1), while interpreting wages and import prices as stationary. In sum, we treat the variables as integrated of order 1 in the econometric modelling.

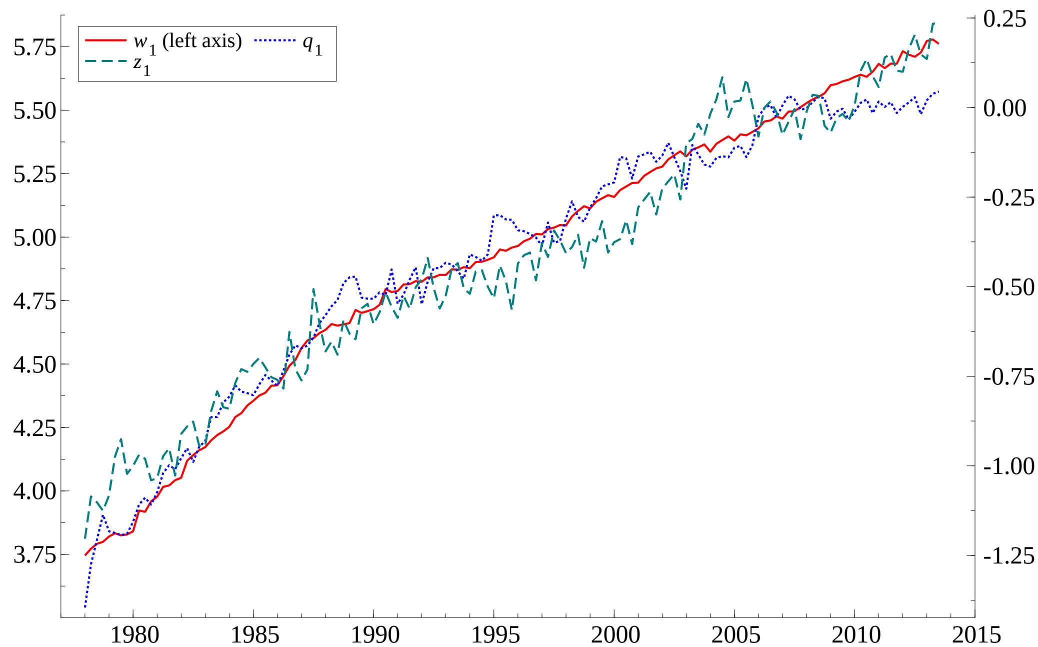

Figure 1: Components of the manufacturing wage share.

Figure 2 shows the unemployment rate, u_t, and the immigration (flow) rate , IM_t. Both variables are percentages (see Appendix A). The unemployment rate and the immigration rate both reject stationarity according to the KPSS test, see Table 1. The graph shows both the first period of departure from post-war full employment in the first half of the 1980s, and the return to near full employment in the years following deregulation of the financial markets (among other things). Next, unemployment increased again during the

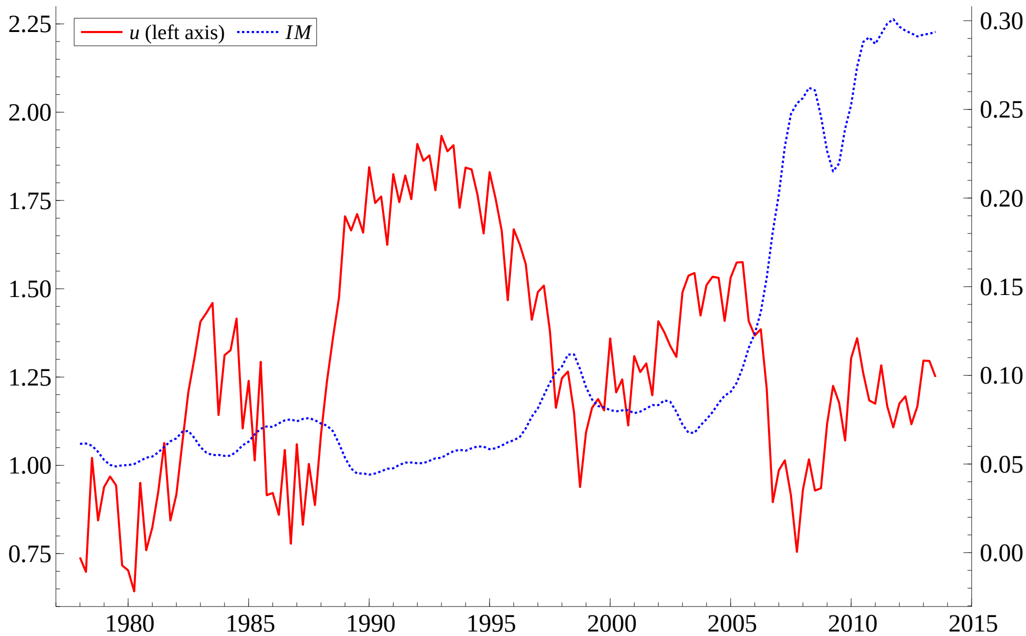

Figure 2: Plot of the logs of the unemployment rate and the immigration flow.

1990s following the fall in oil prices, a restrictive fiscal policy, the collapse of house prices and the ensuing banking crisis. Finally, the small changes in unemployment since 2005 are interesting. The number of employed wage earners in Norway grew by an unprecedented quarter of a million from 2005 to 2009. Naturally, the rate of unemployment fell, but not nearly as much as employment increased. This is due to the high immigration rates. This graph shows labour immigration to Norway from the EU and other developed countries. The migration inflow amounts to 1.6 per cent of the total labour supply in 2011 and is clearly large enough to significantly affect Norwegian labour markets and therefore possibly also wage setting.

We have also investigated the properties of the residuals from a regression of equation (3) and (6), see Table 2. The residuals from regression of equation (3) are labelled e_1, while the residuals from regression of equation (6) are labelled $e_{1,UIM}$. The KPSS test statistic for the wage share in Table 2 is lower and below the critical value if we control for shifts in labour immigration and unemployment. The tests therefore imply that the wage share is stationary when these variables are included. The ADF test does not reject the hypothesis of a unit root, neither for e_1 nor for $e_{1,UIM}$.

3.2 The public sector and private services

We have also investigated the statistical properties in the two other sectors; public and private services. Stationarity is rejected by the KPSS test for the hourly wage rate in both sectors. Figure 3 shows wages in sector 2 relative to sector 1, e_2, together with wages in sector 3 relative to sector 2, e_3. While wages in sector 2 appear to follow wages in sector 1 quite

Table 1: Unit root tests on the variables in levels and first differences. Sample period 1980(1) to 2011(4).[1]

Level				First difference			
Variable	ADF (Lags)	KPSS	Char.[2]	Variable	ADF (Lags)	KPSS	Char.[2]
w_1	-3.714 (1)	0.282	c, t, s	Δw_1	-3.232 (4)	0.878	c, s
z_1	-2.255 (3)	0.076	c, t, s	Δz_1	-10.22 (2)	0.055	c, s
q_1	-2.382 (4)	0.263	c, t, s	Δq_1	-5.550 (3)	0.520	c, s
u	-2.810 (4)	0.205	c, t, s	Δu	-3.841 (3)	0.160	c, s
p	-5.979 (4)	0.310	c, t, s	Δp	-1.842 (3)	1.009	c, s
IM	-0.728 (2)	0.260	c, t, s	ΔIM	-5.352 (1)	0.401	c, s
w_2	-4.303 (4)	0.260	c, t, s	Δw_2	-2.659 (3)	0.818	c, s
w_3	-4.290 (0)	0.192	c, t, s	Δw_3	-9.680 (3)	0.760	c, s

[1]Note, however, that the sample period varies due to the number of lags in the ADF test.

[2]The characteristics are a constant (c), a trend (t) and seasonal dummies (s).

The 5% critical value of the ADF test is -3.45 and -2.88 in level and differences respectively.

The 5% critical value of KPSS test is 0.146 in level and 0.463 in differences.

closely, the wage level in sector 3 was markedly lower than the wage level in sector 2 for a long period. Although this reduction in the relative wage started in late 1980s, it is palusible that it was related to the rise in unemployment in the 1990s and the tight government budgets that characterised the first half of the decade, in particular. This indicates that, in the econometric model, the rate of unemployment can help to explain the development in relative wages. In the last period, wages in sector 3 have increased relative to wages in sector 2, while wages in sector 2 have decreased relative to sector 1. This is the period of high labour immigration.

In Table 2, the residuals from the regression of (7) are labelled e_2, while the residuals from a regression of (7), which includes u and IM, are labelled $e_{2,UI}$. The same regressions are performed for sector 3 as well, and residuals are labelled e_3 and $e_{3,UI}$, respectively. According to the ADF tests in Table 2, e_2, ec_{2UI} and ec_3 are non stationary. Again, however, the KPSS test supports the stationarity of the residuals and indicates that taking account of the possibility that unemployment and immigration can influence the long-run relationships, results in more stationary residuals.

Table 2: Unit root tests on the wage share and the relative wages. Sample period 1980(2) to 2011(4).

Variable	ADF (Lags)	KPSS	Char.[1]
e_1	-2.833 (3)	0.583	c
$e_{1,UIM}$	-2.751 (3)	0.305	c
e_2	-2.439 (3)	0.428	c
$e_{2,UIM}$	-2.232 (3)	0.155	c
e_3	-2.135 (3)	0.328	c
$e_{3,UIM}$	-3.153 (3)	0.170	c

The 5% critical value of the ADF test is -2.88 and 0.463 for the KPSS test
[1]The tests have included a constant (c) in the test.

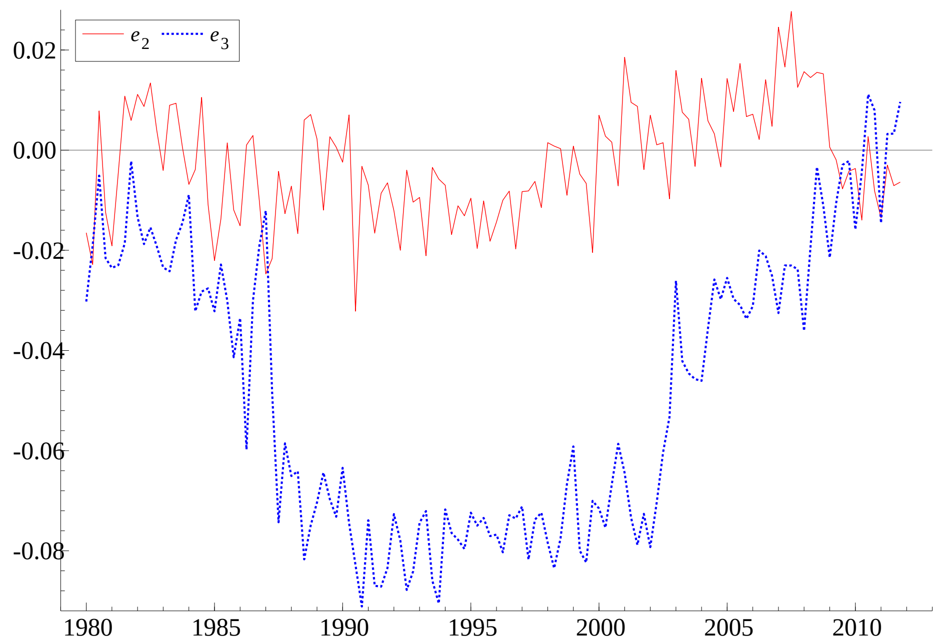

Figure 3: Wages in sector 2 relative to sector 1 (e_2) and wages in sector 3 relative to wages in sector 2 (e_3).

4 Econometric results

The previous section showed that univariate tests give clear support for unit-root non-stationarity, and some initial support for three cointegration relations of the pattern bargaining model. However, cointegration is a system property, and a concise way of testing the long-run theory is to formulate a vector autoregression (VAR) and apply the maximum likelihood method of Johansen (1995b). This step in the analysis is taken in section 4.1.

In section 4.2, we specify a cointegrated equilibrium correction model (ECM) that is consistent with the evidence from the VAR-based test of cointegration. The ECM is used to further test the hypotheses about the dynamics of wage leader- and follower bargaining model (the causal structure between the growth rates). The tests illustrate the relative invariance of the model parameters after the change from a fixed exchange rate to a floating exchange rate and the surge in immigration rates.

4.1 Cointegration

We formulate a fourth order VAR for the three wage levels, manufacturing (w_1), private services (w_2) and public sector (w_3). We include the endogenous variables of the VAR in the vector $\mathbf{Y_{wt}} = (w_{1t}, w_{2t}, w_{3t})'$. The VAR is expressed as:

$$\Delta \mathbf{Y_{wt}} = \sum_{i=1}^{3} \mathbf{\Gamma}_i \Delta \mathbf{Y_{wt-i}} + \mathbf{\Pi} \mathbf{Y_{t-1}} + \mathbf{\Upsilon} \mathbf{Q_t} + \varepsilon_{\mathbf{t}}, \tag{13}$$

where all variables in the cointegration space are defined in the extended data vector: $\mathbf{Y_t} = (\mathbf{Y_{wt}}, x_t, \tau_{1t}, u_t, IM_t)'$. $x = q_1 + z_1$ is the wage-scope variable in the manufacturing sector. The role of the payroll tax rate, τ_{1t}, enables us to distinguish between the wage rate per hour and the wage cost per hour. $\mathbf{\Gamma}_i$ $(i = 1, \ldots, 4)$ and $\mathbf{\Pi}$ are matrices with short-run and long-run parameters for wages, unemployment and immigration.

Formally, τ_{1t}, u_t and IM_t are also treated as I(1), since they are included in the $\mathbf{Y_t}$ vector. However, the tax rate is determined by political decisions and apparent unit root non-stationarity is probably due to policy determined changes in the mean. As the discussion above suggests, while u_t and IM_t are clearly non-stationary, the economically relevant 'model of stationarity' is probably that they are also driven by intermittent structural breaks in their means, rather than being convincingly represented by unit-root non-stationarity.

$\mathbf{Q_t}$ is other non-modelled variables in the VAR, namely the lagged annual CPI inflation rate and dummy variables. For simplicity, they are associated with the single matrix of parameters $\mathbf{\Upsilon}$. Finally, the vector $\varepsilon_{\mathbf{t}}$ contains the VAR residuals for the three wage growth rates.

As noted above, $\mathbf{Q_t}$ in (13) implies that the VAR includes other variables as well. In terms of economic interpretation, the most important variable is the lagged rate of consumer price inflation, which is included to control for the effects of cost of living considerations (as noted above). In the sample period, there have been periods of centralised wage settlements (wage laws), see e.g. Bowitz and Cappelen (2001) and Section 3. We include dummies for these events, and for other quarters with outliers. Finally, we include a constant, three centred seasonal dummies and a deterministic trend. The result is a relatively well-specified VAR. The residuals, $\hat{\varepsilon}_{\mathbf{t}}$, have a tendency to be negatively auto-correlated, indicating a certain over-fitting. However, the test of joint non-normality of $\varepsilon_{\mathbf{t}}$ is insignificant at the 5 % significance level.

$\mathbf{\Pi}$ is often referred to as the long-run matrix, and it has the dimensions 3×7. Let r denote the rank of $\mathbf{\Pi}$. Since x_t, τ_{1t}, u and IM are non-modelled variables, r can be 0, 1, 2 or 3. $r = 0$ means that there is no cointegration, which would imply rejection of the theory we

formulated above. If, on the other hand, $r = 3$, the variables x_t, τ_{1t}, u and IM represents four common trends in the three nominal wage rates. The theory of wage leadership points to the wage-scope variable $x_t = q_{1t} + z_{1t}$ as the most dominant trend in the system. This is because the evolution of the scope variable defines the upward trend in the wage corridor in the manufacturing sector. If wage leadership holds, this trend is transmitted to the two other wage rates.

$\mathbf{\Pi}$ can be written as $\mathbf{\Pi} = \boldsymbol{\alpha\beta}'$, where $\boldsymbol{\alpha}$ is $3 \times r$ and $\boldsymbol{\beta}$ ($7 \times r$) is the matrix with cointegration parameters. The theory presented in section 2 implies $\tau = 3$ with (3), (7) and (10) as the cointegration relationship.

When conducting the cointegration rank test a trend is included and restricted to lie in the cointegration space (i.e. included in $\mathbf{Y_{t-1}}$) such that critical value exist, see Harbo et al. (1998). Table 3, column *Trace test statistic* presents the test statistic of the cointegration analysis. The critical values of Doornik (2003) for zero, one and two cointegrating relationships are presented in Table 3, column five. The trace test statistic for two or fewer cointegration relationships is below the 10 % critical value for a standard integrated VAR reported in Doornik (2003). This implies that the test rejects two or fewer cointegration relationships at the 10 % significance level. This result supports the hypothesis of three cointegrating relationships. However, these critical values do not take into account exogenous variables. The critical values increase with the number of exogenous variables included, cf. Harbo et al. (1998). The critical values with four exogenous variables are shown in Table 3, last column.

The two first test statistics are well above their respective critical values (adjusted for exogenous variables), which supports that the number of stationary long-run relationships is at least two. The last row shows that we are unable to formally reject two, and accept three, cointegrating vectors. However, the critical values are for asymptotic distributions, and using these critical values may lead to over-rejecting in small samples, see Doornik (1998). The correction factor for critical values without exogenous variables is given in Johansen (2002). Correction factors do not exist for critical values in cointegration analyses with exogenous variables, but the correction factor without exogenous factors implies that the critical values of Harbo et al. (1998) should be interpreted with care. We conclude that cointegration tests, broadly interpreted, provide formal support for three cointegrating relationships ($r = 3$), but that equilibrium correction may be relatively weak.

Table 4, Panel 1, shows the maximised likelihood value, *Log L*, after we have omitted the deterministic trend from the three long-run relationships. The trend was included for the purpose of making correct inferences about the rank, but once the rank has been fixed to three, its significance can be tested by a standard Likelihood-Ratio (LR) test with a Chi-square distribution. In this case, the restriction is rejected at the 1 % level. However, keeping a trend is without meaning economically, since the wage share and relative wages will then drift deterministically. A closer inspection of the results also reveal that the trend is sensitive to if the first three years in public sector is included. There was a substantial wage stop in public sector which ended in the beginning of the 80s and wages increased quite rapidly in the following years. Hence, if the first three years are excluded the tests imply that the trend should be removed. We therefore conclude that the trend should be removed from the cointegration space. In Table 4, we use the result of a rank equal to 3, but use a VAR without trend when testing the identification of the cointegrating relationships.

Table 3: Tests of cointegration rank. Sample 1982(1)-2011(4).

Eigenvalue			*Trace*	10% *Critical value Doornik*	
(λ_i)	H_0	H_1	*Test statistic*	*No exogenous*	*With exogenous*
0.37	$r = 0$	$r >= 1$	116.5	39.7	60.5
0.23	$r <= 1$	$r >= 2$	51.3	23.3	37.7
0.17	$r <= 2$	$r >= 3$	14.3	10.7	18.5

Endogenous variables: wages in manufacturing (w_1), private services (w_2) and public sector (w_3).
Restricted variables: wage scope ($x = q_1 + z_1$) in manufacturing,
unemployment (u), immigration (IM) and a trend (t).
Unrestriced variables: payroll tax-rate, τ_1, constant, seasonal dummies and some specific dummies which capture important events in the Norwegian wage setting, see e.g. Bowitz and Cappelen (2001) and Section 4.

Removing the deterministic trend does not contribute to identification. In Panel 1, we therefore show the results after we have imposed identifying restrictions that are consistent with the wage leadership theory. Specifically, the payroll tax rate is only included in the cointegrating relationship, which is normalised on w_1. This separates the second and third relationships from the first. The exclusion of x from the second and third relationships provides more identifying information for these two relationships. In order to separate the second from the third relationship, we exclude w_1 and IM from the third, and w_3 from the second relationship. Finally, in order to identify the first relationship, both w_2 and w_3 are excluded from the relationship that is normalised on w_1. The identified long-run relationships are reported in Panel 1 of the table, together with the likelihood value and the Chi-square distributed test of the seven over-identifying restrictions that represent the wage leadership view of Norwegian wage formation. The restrictions are statistically acceptable at the 5 % level.

Table 4: Testing steady-state hypotheses

Unrestricted system for wages

$w_1 = \beta_{w2,1}w_2 + \beta_{w3,1}w_3 + \beta_{\tau,1}\tau + \beta_{x,1}x + \beta_{u,1}u + \beta_{IM,1}IM$

$w_2 = \beta_{w1,2}w_1 + \beta_{w3,2}w_3 + \beta_{\tau,3}\tau + \beta_{x,2}x + \beta_{u,2}u + \beta_{IM,2}IM$

$w_3 = \beta_{w1,3}w_1 + \beta_{w2,3}w_2 + \beta_{\tau,3}\tau + \beta_{x,3}x + \beta_{u,3}u + \beta_{IM,3}IM$

Unrestricted $Log\ L = 1365.79768$

Panel 1: Identified long-run relationships

$w_1 = -\underset{(.)}{1}\,\tau + \underset{(.)}{1}\,x \underset{(0.09)}{-0.38}\,u \underset{(0.42)}{-1.72}\,IM$

$w_2 = \underset{(.)}{1}\,w_1 \underset{(0.07)}{-0.40}\,u \underset{(0.20)}{-0.51}\,IM$

$w_3 = \underset{(.)}{1}\,w_2 \underset{(0.014)}{-0.07}\,u$

$Log\ L = 1359.27709$, $\chi^2(7) = 13.04[0.07]$

Panel 2: Manufacturing wage leadership, $\alpha_{w1,2} = \alpha_{w1,3} = 0$

$w_1 = -\underset{(.)}{1}\,\tau + \underset{(.)}{1}\,x \underset{(0.09)}{-0.41}\,u \underset{(0.43)}{-1.76}\,IM$

$w_2 = \underset{(.)}{1}\,w_1 \underset{(0.07)}{-0.40}\,u \underset{(0.21)}{-0.57}\,IM$

$w_3 = \underset{(.)}{1}\,w_2 \underset{(0.015)}{-0.08}\,u$

$Log\ L = 1359.25748$, $\chi^2(9) = 13.08[0.16]$

Additional restrictions: $\chi^2(2) = 0.04(0.98)$

Panel 3: No response of Δw_2 to $(w_{2t-1} - w_{3t-1})$, $\alpha_{w2,3} = 0$

$w_1 = -\underset{(.)}{1}\,\tau + \underset{(.)}{1}\,x \underset{(0.09)}{-0.41}\,u \underset{(0.43)}{-1.76}\,IM$

$w_2 = \underset{(.)}{1}\,w_1 \underset{(0.07)}{-0.42}\,u \underset{(0.19)}{-0.55}\,IM$

$w_3 = \underset{(.)}{1}\,w_2 \underset{(0.016)}{-0.08}\,u$

$Log\ L = 1359.18919$, $\chi^2(10) = 13.22[0.21]$

Additional restrictions: $\chi^2(1) = 0.14[0.71]$

Panel 4: No direct response of Δw_3 to manufacturing norm, $\alpha_{w3,1} = 0$

$w_1 = -\underset{(.)}{1}\,\tau + \underset{(.)}{1}\,x \underset{(0.12)}{-0.58}\,u \underset{(0.53)}{-2.45}\,IM$

$w_2 = \underset{(.)}{1}\,w_1 \underset{(0.03)}{-0.17}\,u \underset{(0.15)}{-0.18}\,IM$

$w_3 = \underset{(.)}{1}\,w_2 \underset{(0.01)}{-0.04}\,u$

$Log\ L = 1358.39232$, $\chi^2(11) = 14.81[0.19]$

Additional restrictions: $\chi^2(1) = 1.59[0.21]$

Conditional on the identified long-run relationships, we can begin to test for the minimum restrictions that imply no feed-back from w_2 and w_3 to w_1. The results are reported in Panel 2 of the table, using $\alpha_{i,j}$ to denote the element in row i, column j in the matrix with loadings parameters. The increase in the test of the over-identifying restrictions is small. Therefore it is not surprising that two zero restrictions that distinguish Panel 2 from Panel 1, are statistically acceptable (the incremental test yields $\chi^2(2) = 0.04[0.98]$). We note that the estimated cointegration parameters also change very little from Panel 1 to Panel 2.

We conclude that the tests in Panel 1 and 2 provide relatively clear support for the view that the wage-leading role of the manufacturing sector has been in operation during our sample period. The result is different from those estimated by Nymoen (1991), who was unable to impose the wage-leading restrictions when estimating on a sample that started in 1969(1) and ended in 1987(4). As a result, he specified a preferred model that allowed for wage-wage effects. The different samples might be relevant here, since the 1970s in particular were characterized by super high wage inflation in certain years. Some of this may have been driven by low coordination and wage-wage inflation.[5]

Returning to Table 4, Panel 3 and 4 test additional restrictions implied by pattern wage bargaining. In panel 3, we add the restriction that the deviations from the second and third cointegration relationships are not corrected by manufacturing wages. This result shows that the increase in the test of over-identifying restrictions is small, and the incremental test does not reject additional restrictions. The cointegration parameters are also very similar when we compare Panel 2 and Panel 3. Finally, in Panel 4, the restriction that wages in sector 3 does not equilibrium correct with respect to the first cointegrating relationship is imposed. Although this restriction is acceptable, we note that the parameter estimates change notably. We return to the question of a possible contemporaneous effect of w_1 on w_3 in the structural model below.

The point estimates in the best identified models, e.g. in Panel 3, deserve comment. In the manufacturing sector (Sector 1), the elasticity of wages ($\beta_{u,1}$) with respect to unemployment is higher in absolute value than the estimates in, e.g., Nymoen (1989) and Johansen (1995a), who used data from the 1980s.

This may be due to the larger variation of u_t in our sample, which includes the period after the Norwegian banking crisis in the 1990s. For Sector 2, the estimated unemployment elasticity is practically the same ($\hat{\beta}_{u,2} = -0.4$). We note, however, that this implies that the estimated long-run wage responsiveness for the private wage-following sector is twice as large as for the wage-leading sector. This is realistic, since this sector is characterized by weaker union organisation and lower capacity for multi-employer agreements than in the manufacturing sector. In the public sector, the estimated parameter of unemployment ($\beta_{u,3}$) is much lower (in absolute value) than in the two other sectors, which is also reasonable given that changes in overall unemployment have little direct relevance for the two bargaining parties.

Since labour immigration is a new phenomenon, there are no studies that can serve as direct references for discussing our results for the parameters $\beta_{IM,1}$ and $\beta_{IM,2}$ in Panel 3.

[5]There are other differences as well, both in the data definitions and in the econometric methodology. Nymoen did not identify the cointegrating relationships separately from the short-run dynamics, as we will do. This may have led to poorer identification of his long-run relationships.

However, one important study that use micro data reports evidence that immigration has causally reduced wage earnings in certain sectors of the Norwegian economy, see Bratsberg and Raaum (2012). Compared to our own earlier research using single equation estimation methods, the estimated value of $\hat{\beta}_{IM,1} = -1.7$ is very large, see Gjelsvik et al. (2013). Note, however, that the estimated standard error indicates a broad 95 % confidence interval. The single equation estimation just mentioned provides point estimates of somewhat less than -1 for the wage-leading sector, and this value is within the confidence interval with good margin.

Based on the above interpretation of the evidence we have specified the following equilibrium correction variables for use in a structural econometric model:

$$\widehat{ec}_{1t} = w_{1t} + \tau_t - x_t + 0.38u_t + 1.1IM_t \tag{14}$$

$$\widehat{ec}_{2t} = w_{2t} - w_{1t} + 0.40u_t + 0.51IM_t \tag{15}$$

$$\widehat{ec}_{3t} = w_{3t} - w_{2t} + 0.07u_t \tag{16}$$

The only significant difference from Panel 1, is that we have adjusted the point estimate of the coefficient of IM in (14). The estimation results of the dynamic structural model will confirm that this downsizing of the direct effect of immigration on the manufacturing wage is data-acceptable.

4.2 A dynamic simultaneous equations econometric model

In this section, we formulate and estimate a dynamic model model that is consistent with the cointegration analysis, and which is identified econometrically. At this stage in the analysis we include CPI inflation into the model in an explicit way, which was suggested to be important to wage formation in Section 2.

In the VAR above, the lagged four-quarter change in the log of the official consumer price index, denoted $\Delta_4 p_{t-1}$ was included among the variables in $\mathbf{Q_t}$ in (13). In the same way as for the lagged wage growth rates, we can think of lagged price inflation as a variable in a reduced form obtained from a simultaneous equations model, SEM, that includes price inflation among the endogenous variables. In order to model price growth jointly with wage growth, we need a separate long-run relationship for the price level. We do not carry out a cointegration analysis for the price level in this paper. Instead, we use the following equilibrium correction term

$$\widehat{ec}_{4t} = p_t - 0.6(w_{1t} + \tau_t - z_{1t}) - 0.4pi_i \tag{17}$$

where the elasticities for unit wage cost and import price, pi, are taken from the econometric results in Bårdsen et al. (2003).

The SEM model can formally be expressed as:

$$\mathbf{B_0 Y_{wpt}} = \sum_{i=1}^{3} \mathbf{B}_i \mathbf{Y_{wpt-i}} + \mathbf{AEC_{t-1}} + \Upsilon \mathbf{Q_t} + \epsilon_\mathbf{t}, \tag{18}$$

where $\mathbf{Y_{wpt}} = (\Delta w_{1t},\ \Delta w_{2t},\ \Delta w_{3t}, \Delta_4 p_t)'$ contains changes in wages and the consumer price index, $\mathbf{EC}$ contains the long run relationships in (14)-(17) and $\mathbf{Q_t}$ is redefined to exclude

consumer prices. $\mathbf{A}$ is a diagonal matrix and all the elements of the diagonal in $\mathbf{B_0}$ is equal to one.

The last four columns in table 5 show coefficient estimates ($\hat{\mathbf{A}}$) for an identified four equation SEM, where a diagonal matrix with adjustment coefficients for the four ec-terms is sufficient for identification on the rank condition.[6] The maximised log likelihood reported in the last row serves as the unrestricted likelihood value that we use to test the wage-leadership-followership restrictions in Table 6.[7] The two test statistics are (F_{AR}), for autocorrelation of order between 1 and 5, and non-normality (χ^2_{NORM}), calculated from the SEM residuals. They are system versions of autocorrelation and normality tests, see Doornik and Hendry (2013), and are reported with asymptotic p-values in square brackets.

Table 5: Wage-price SEM. FIML estimates of the contemporaneous coefficient matrix $\hat{\mathbf{B}}_\mathbf{0}$ (first four columns with estimates) and adjustment coefficients $\hat{\mathbf{A}}$ (last four columns). Standard errors in brackets below the estimates. Sample period 1980(1) to 2011(4).

	Δw_{1t}	Δw_{2t}	Δw_{3t}	$\Delta_4 p_t$	$\widehat{ec}_{1t-1}$	$\widehat{ec}_{2t-1}$	$\widehat{ec}_{3t-1}$	$\widehat{ec}_{4t-1}$
Δw_{1t}	-1	-0.05 (0.15)	-0.04 (0.10)	0.37 (0.06)	-0.04 (0.01)	0	0	0
Δw_{2t}	0.42 (0.06)	-1	0.03 (0.06)	0.05 (0.09)	0	-0.04 (0.006)	0	0
Δw_{3t}	0.17 (0.08)	0.24 (0.12)	-1	0.16 (0.03)	0	0	-0.19 (0.03)	0
$\Delta_4 p_t$	0.02 (0.06)	-0.12 (0.12)	0.09 (0.09)	-1	0	0	0	-0.10 (0.02)

$Log\ L$= 1892.0816, $F_{AR} = 1.18[0.16]$ and $\chi^2_{NORM} = 12.6[0.12]$.

Table 5 shows that the diagonal elements of the matrix with adjustments coefficients for the ec-terms are statistically significant, although the numerical values are relatively small for the first two. This is consistent with the results in Table 3 and Table 4. The results for the contemporaneous parameters ($\hat{\mathbf{B}}_\mathbf{0}$) also show a clear pattern. In the row for Δw_{1t}, the two coefficients of Δw_{2t} and Δw_{3t} are negative ("wrong sign"), but they are statistically insignificant from zero values (judged by the standard errors). Conversely, Δw_{1t} has a sizeable estimated coefficient in the rows for Δw_{2t} and Δw_{3t}, which supports the hypothesis that the manufacturing sector is wage-leading, with private service production and the public sector as wage-followers.

The first row in Table 6 shows that three versions of the null hypothesis that manufacturing is *not* wage-leading, are formally rejected at the 5 % level. In the second line of tests, the second entry supports that the private service sector is wage-leading in relation to the public sector. The third row of tests shows that the coefficient of Δw_{3t} can be restricted to

[6]At this stage, there are no restrictions on the correlation matrix of the contemporaneous disturbances, which means that the rank condition is necessary and sufficient.

[7]Note that this likelihood is not comparable to the corresponding in Table 4, since the model in Table 5 includes CPI inflation as an endogenous variable

Table 6: Likelihood-ratio tests of wage-leader/follower restrictions on the model in Table 5

Restrictions:	Sec 1 $\nrightarrow$ Sec 2 $\chi^2(1) = 54.5^{**}$	Sec 1 $\nrightarrow$ Sec 3 $\chi^2(1) = 4.6^{*}$	Sec1 $\nrightarrow$ Sec 2 and 3 $\chi^2(2) = 54.9^{**}$
Restrictions:	Sec 2 $\nrightarrow$ Sec 1 $\chi^2(1) = 0.11$	Sec 2 $\nrightarrow$ Sec 3 $\chi^2(1) = 4.4^{*}$	Sec 2 $\nrightarrow$ Sec 1 and 3 $\chi^2(2) = 4.7$
Restrictions:	Sec 3 $\nrightarrow$ Sec 1 $\chi^2(1) = 0.19$	Sec 3 $\nrightarrow$ Sec 2 $\chi^2(1) = 0.20$	Sec 3 $\nrightarrow$ Sec 1 and 2 $\chi^2(2) = 0.42$

* and **denotes significance at the 5% and 1 % levels.

zero in both the manufacturing sector wage equation and in the private service sector wage equation without any significant drop in the likelihood value.

We next turn to the role of price inflation as measured by $\Delta_4 p_t$. This variable is significant with a large estimated coefficient in the manufacturing wage equation. The coefficient of inflation in the equation for Δw_{2t} is close to zero, but can note that there is a significant effect of the lagged inflation rate in this equation, see Appendix B with detailed estimation results.

The last row of Table 5, with the results for the inflation equation $\Delta_4 p_t$, shows that there is little support for within-quarter effects of wage changes on inflation.[8] However, since $\widehat{ec}_{4t}$ includes wage costs (w_{1t}+ $\ln(1+\tau_t)$), we nevertheless have a closed feed-back loop between wages and CPI.

Table 7 shows the results of a restricted estimation where we have imposed all of the six exclusion restrictions discussed above. Compared to Table 5, the retained coefficients change very little.

Table 7: Wage-price SEM with wage leader-follower restrictions imposed. FIML estimates of $\hat{\mathbf{B}}_{\mathbf{0}}$ and $\hat{\mathbf{A}}$ in equation (18) with standard errors in brackets below the estimates. Sample period 1980(1) to 2011(4).

	Δw_{1t}	Δw_{2t}	Δw_{3t}	$\Delta_4 p_t$	$\widehat{ec}_{1t-1}$	$\widehat{ec}_{2t-1}$	$\widehat{ec}_{3t-1}$	$\widehat{ec}_{4t-1}$
Δw_{1t}	-1	0	0	0.35 (0.06)	-0.04 (0.01)	0	0	0
Δw_{2t}	0.42 (0.06)	-1	0	0.06 (0.09)	0	-0.04 (0.01)	0	0
Δw_{3t}	0.17 (0.08)	0.25 (0.12)	-1	0.16 (0.03)	0	0	-0.19 (0.03)	0
$\Delta_4 p_t$	0	0	0	-1	0	0	0	-0.10 (0.02)

Log L =1890.98642, $F_{AR} = 1.17[0.17]$ and $\chi^2_{NORM} = 13.0[0.11]$
LR test of wage-leader following restrictions: $\chi^2(6) = 2.19[0.14]$

[8] The estimated SEM includes three unrestricted lags in the quarterly inflation rate in the fourth row. This ensures that there are no restrictions implied by normalisation on $\Delta_4 p_t$ in the fourth row. This representation is convenient for modelling the effects of inflation on wage dynamics.

It is interesting that Table 5 and Table 7 show that inflation expectations[9] have a significant effect on wage formation in the manufacturing sector, and therefore on the wage norm of the system. Is this evidence of inflation targeting in the sense that the wage increases in the manufacturing sector are led by the inflation forecast of the Central Bank rather than by the evolution of the wage scope? In order to investigate this possibility, as well as the invariance of the system with respect to the monetary policy regime shift more generally, we turn to the estimation results on a sample that ends before the structural break in the market for foreign exchange and in interest rate determination.

4.3 Stability and invariance

In order to address the the question of stability of the wage formation structure, we have estimated the model on a sample that ends in 2000(4). This sample ends before the labour immigration from Europe started, and it is unlikely that the immigration effects on wages that we estimated above are invariant to the shortening of the sample.

Consequently, it does not make sense to define the three wage equilibrium correction terms in the same way as in (14)-(16). Instead, we remove IM_t and u_t from the equilibrium correction variables, which for the short sample are defined as:

$$\widehat{ecs}_{1t} = w_{1t} + \tau_t - x_{1t} \tag{19}$$

$$\widehat{ecs}_{2t} = w_{2t} - w_{1t} \tag{20}$$

$$\widehat{ecs}_{3t} = w_{3t} - w_{2t} \tag{21}$$

implying the same long-run relationships as in the full sample. The estimated system is therefore specified by including immigration and unemployment as exogenous I(0)-variables (but subject to breaks) in the system. For the CPI level, we use (17) as the equilibrium correction term also in the short sample.

The results of the estimation with the wage-leadership-follower restrictions imposed are shown in Table 8. The estimation results are very similar to the full sample estimates in Table 7, although the joint test of the six restrictions is marginally significant at the 5 % level. However, and as previously discussed, the test statistic is not corrected for the smaller sample size. The details show that this is mainly due to an estimated negative coefficient of Δw_{2t} in the Δw_{1t} equation, i.e. the same sign problem that we noted for the full sample results, but more pronounced on the short sample. Appendix B contains Table 10 with the detailed dynamic specification for the short sample.

[9]Inflation in period t is included as an endogenous variable in the model estimated by FIML. Therefore, the inflation variable in the manufacturing sector wage equation for example, can be interpreted as an expectations variable.

Table 8: Pre inflation targeting wage-price SEM with wage-leader-follower restrictions imposed. FIML estimates with standard errors in round brackets below the estimates. Sample period 1980(1) to 2000(4).

	Δw_{1t}	Δw_{2t}	Δw_{3t}	$\Delta_4 p_t$	$\widehat{ecs}_{1t-1}$	$\widehat{ecs}_{2t-1}$	$\widehat{ecs}_{3t-1}$	$\widehat{ec}_{4t-1}$
Δw_{1t}	-1	0	0	0.32 (0.06)	-0.05 (0.02)	0	0	0
Δw_{2t}	0.45 (0.06)	-1	0	0.11 (0.13)	0	-0.11 (0.09)	0	0
Δw_{3t}	0.15 (0.09)	0.49 (0.14)	-1	0.16 (0.04)	0	0	-0.24 (0.08)	0
$\Delta_4 p_t$	0	0	0	-1	0	0	0	-0.08 (0.02)

Log L =1288.14658, $F_{AR} = 1.10[0.29]$ and $\chi^2_{NORM} = 14.3[0.08]$
LR test of wage-leader following restrictions: $\chi^2(6) = 12.39[0.05]$

The results show in particular that in manufacturing the estimated coefficient of $\Delta_4 p_t$ is 0.32, with a t-value of 5.1. This is very close to the full sample results. Taken together, the two estimates show that the effect of inflation expectations on the wage norm was in place before inflation targeting was introduced. This is not surprising since CPI expectations have been an important part of the wage setting system for decades. Coordination of inflation expectations among representatives of the employer and employee confederations has been one of the main purposes of the TCC institution long before inflation targeting was introduced.

Another way of illustrating the stability of the estimated structure is by forecasting. Figure 4 shows dynamic forecasts for the period 2001(1) to 2011(4). As can be expected, there are some examples of forecast failures (actuals outside the forecast uncertainty fans), in particular for inflation. However, a major structural break in the 40-quarter forecast period would have resulted in much clearer forecast.

Finally, we illustrate relative parameter stability in the dynamic multipliers of the structural model. In Figure 5, the impulse responses to a shock in the inflation equation (an inflation shock) are shown. There are two graphs in each panel, corresponding to the short sample estimation and the full sample estimation. In all panels, the dynamic multipliers are very similar, illustrating that the results of this policy analysis do not depend on the ten years of inflation targeting. The sign of the multipliers changes from positive to negative. The interpretation is that, because the long-run growth rate in the manufacturing sector depends on producer price growth and productivity growth, the short-run influence of a shock to inflation will mean overshooting in manufacturing sector wages. Because of the wage-leader role of manufacturing, the overshooting spills over to wages in the two other sectors.

Figure 6 shows the impulse response parameters of a shock to the immigration rate (IM). These impulse responses are not invariant. As noted above, this is not surprising given that the surge in immigration came after 2000 (cf. Figure 2). Specifically, it is the effects of increased immigration on wage formation in the manufacturing sector that are underestimated on the sample that ends in 2000(4). Since wage growth in the wage-leading sector is important for the overall nominal path of the Norwegian economy, this spills over

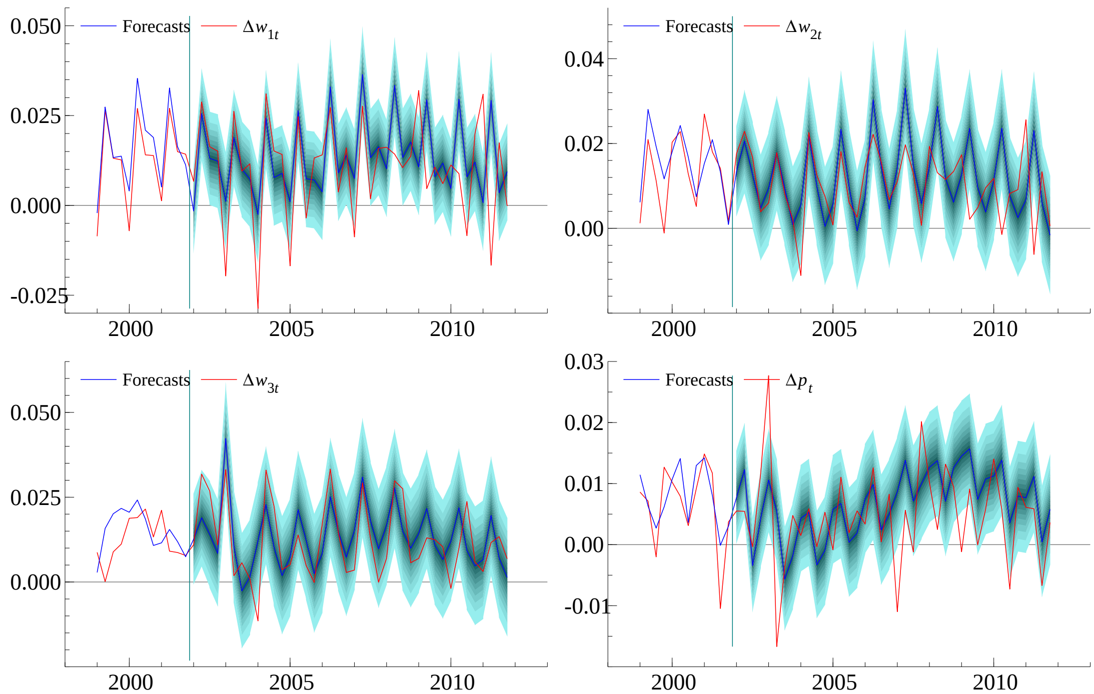

Figure 4: Dynamic forecasts with a 95 per cent forecast uncertainty fans. Forecasts are in blue, actuals in red. The estimation period is from 1980(1) to 2000(4).

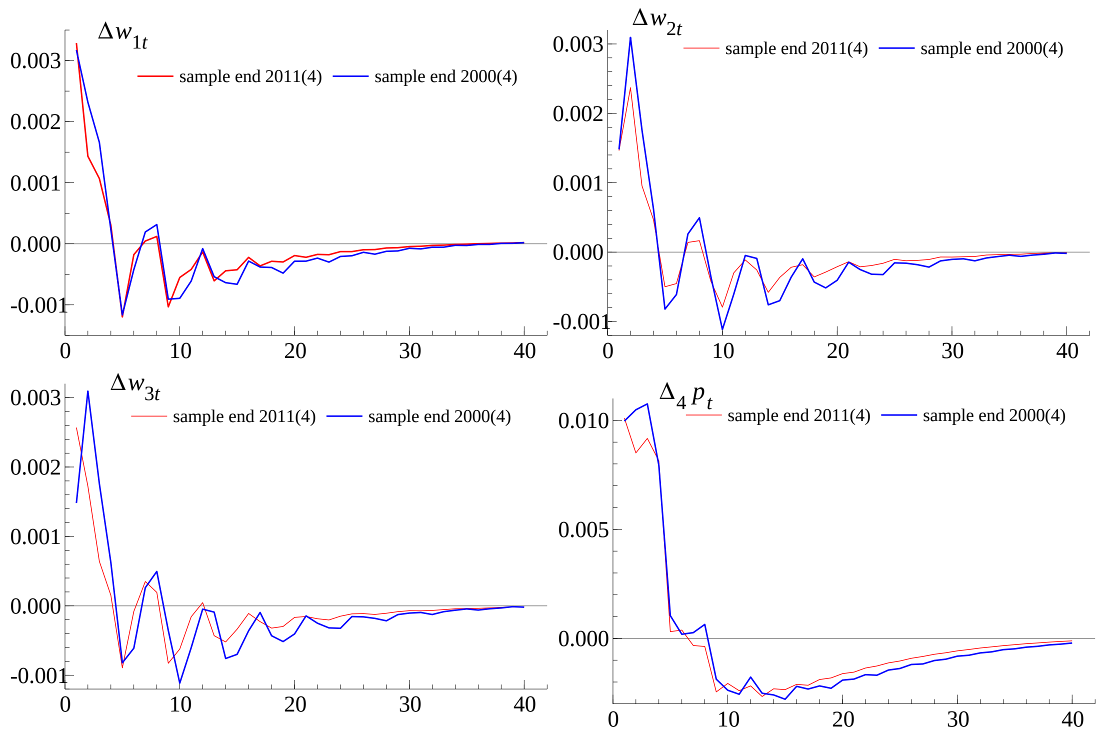

Figure 5: Impulse responses of a shock to the rate of quarterly inflation of 0.01 per cent.

to the inflation impulse responses in particular, which are severely underestimated on the sample that ends in 2000(4).

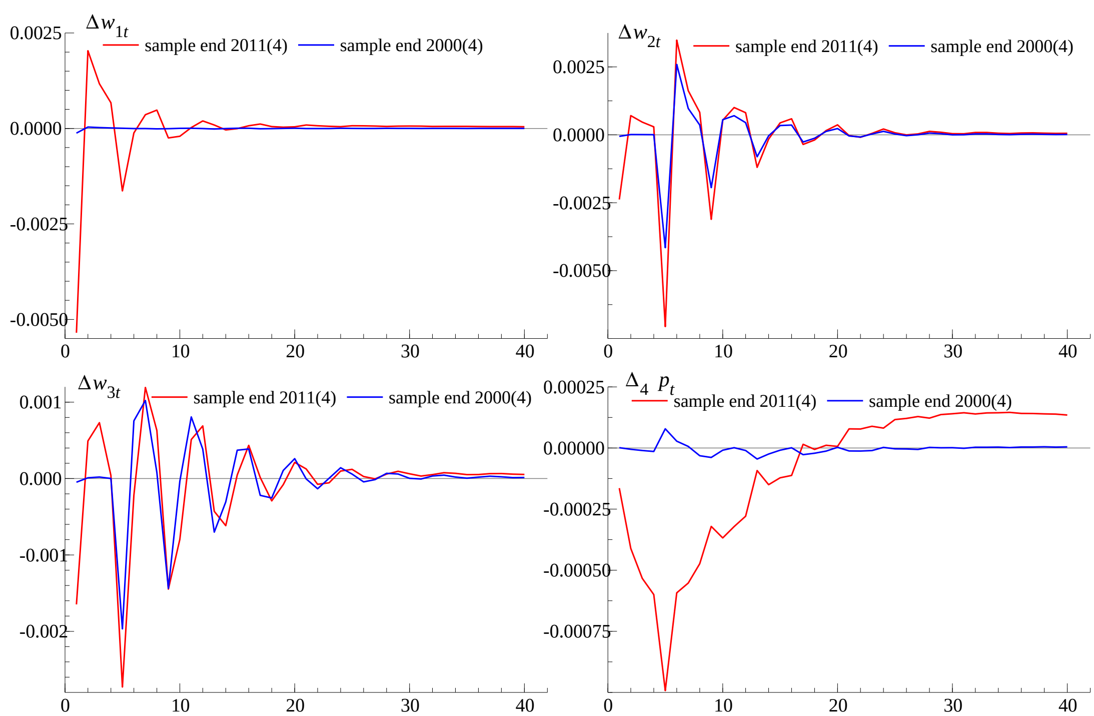

Figure 6: Impulse responses of a shock to the immigration rate by 0.10 percentage points.

5 Conclusion

In this paper we have modelled wages in three sectors of the Norwegian economy; manufacturing, private services and public sector. The model represents the typical pattern of wage bargaining in Norway, where manufacturing sector negotiates first and where the outcome is considered to be the wage norm for the other sectors of the economy. The logic of this pattern is simple, if the wage growth in the exposed sector becomes a wage norm for wage setting in the other sectors, the functional income distribution will be relative both in the wage-leading and wage-following sectors. The wage norm is defined in a system of arbitration that involves self-interest among both workers and firms, but, fundamentally also compromise. As noted in the introduction, the analysis of the system of wage formation as a 'social institution' applies quite, cf Solow (1990). generally. What is perhaps more special about Norway is the continued importance of collective wage agreements, and the roles that the confederate labour market organizations hold the tripartite interaction between government and social partners.[10].

Over the last 10-15 years, there have been two specific challenges to the Norwegian system of wage formation. First, the change to a floating exchange rate and, as a consequence, inflation targeting monetary policy. Second, the rise in labour immigration from outside Scandinavia. Both changes could potentially affect the bargaining power or the gain of coordinating wage formation, and hence change the recursive structure of wage formation in Norway.

The analytical framework that we specify, allows us to test the historical pattern of wage bargaining with a model with no pattern imposed. The model also allows the level of net immigration flow to affect wages separately.

The empirical results give the following answers to the questions posed in the introduction. The cointegration analysis show that there is a stable long term relationship between wage levels in the three sectors, with the manufacturing sector as the wage leader. Therefore, wage growth in manufacturing still constitute the wage norm of the Norwegian wage formation. The empirical results show that the wage norm in manufacturing depends heavily on the profitability growth in the same sector, which is consistent with the theory of collective wage bargaining, but the long term wage level is lower due to the negative and significant effect of the immigration flow. Immigration has also affected the relative wage between private and public services. This might suggest that immigration has affected bargaining power (a parameter of the system), but so far without fundamentally altering the pattern wage bargaining system. And although both unions and employment confederations have understood the nature and aims of monetary policy, we find no indication that the Norwegian monetary policy in 2001 has changed the system of wage formation or that inflation expectations have become more important.

Immigration effects apart, we find that the econometric model has more or less the same parameter estimates on two different samples: one with the period of inflation targeting, and another that ends before this policy change was formally introduced in 2001. The impulse responses of a shock to inflation also show a high degree of invariance. The recursive structure where the wage settlements in manufacturing represent a norm for wage setting is still an

[10]cf. OECD (2012, p. 15)

important part of the Norwegian wage settlements, despite both the structural break in monetary policy and high immigration rates. Hence, the corrosion of, or the attacks on the system of coordinated bargaining analyzed by Marginson (2014) have not been seen in Norway. Instead, the level of coordination is probably notably higher than it was during the 1980s, as also Visser (2013) concludes in his broad analysis of trends in wage bargaining institutions , see Visser (2013, p. 63)

One of the explanations why inflation targeting had little impact on the structure of wage setting, may be that the system of wage formation already included an important element of inflation forecasting and expectations formation. Formally, the Technical Calculation Committee for Wage Settlements has for decades helped to synchronize the inflation anticipations of unions and by firms. For this reason, the central bank's focus on inflation forecast and expectations when inflation targeting started in 2001, was not as "new" to wage and price setters as some commentators would have it.

There are of course other aspects of monetary policy that affect how difficult or easy it is to reconcile inflation targeting with the Norwegian system of wage formation. It is quite likely that a practice of strict inflation targeting (large weight on the inflation-gap and short policy horizon) would have put stress on the system. However, Norges Bank has presented itself as a super flexible inflation targeter. The indication is that the prospects for both unemployment and GDP growth carried large weights in the interest rate policy decisions already before the financial crisis in 2008. During most of the period of inflation targeting, Norges Bank have adopted a relatively long policy horizon, and has shown no haste in closing inflation gaps. The possibility that monetary policy in the future will become more strict to changes in wage setting remains only a vague prospect.

A Data definitions and sources

As explained in the text, lower case letters refer to the logarithm of the original variables listed below. For example, $u_t = \log(U_t)$ denotes the log of the unemployment rate. Variables in first differences are denoted by Δ. Subscripts denote time period. For example, p_{t-4} refers to the (log of) the price level four periods back.

$W_{it}-$Index for hourly wage in sector i=1,2,3
P_t- Consumer price index
$Q_{1t}-$ Price deflator of gross value added, manufacturing industry
$Z_{1t}-$ Labour productivity, output per hour in manufacturing
PI_t- Price deflator on imports of goods and services
U_t- Unemployment rate, in per cent. Civilian unemployment,
IM_t- Immigration from land group 1 and 2, in percent of the population aged 15-74. Group 1 include EU/EFTA countries, North America, Australia and New Zealand. Group 2 includes Eastern Europe except EU countries.
τ_t- represents the natural logarithm of the payroll tax rate plus one.

All variables are from the database of the macroeconometric model KVARTS, maintained by Statistics Norway.

B Additional estimation results

In this appendix, Table 9 shows the detailed dynamic specification of the model in Table 7, while Table 10 is the counterpart to Table 8 in the main text.

Table 9: The dynamic specification of the leader-followership model in Table 7

Sector 1 (manufacturing)
$\Delta w_{1t} = -\underset{(0.0368)}{0.1198}\ \Delta\ln(1+\tau_t) - \underset{(0.0489)}{0.3834}\ \Delta_3 w_{1t-1} - \underset{(0.00797)}{0.04055}\ \widehat{ec}_{1t-1} + \underset{(0.042)}{0.3462}\ \Delta_4 p_t$
Sector 2 (private service)
$\Delta w_{2t} = \underset{(0.0558)}{0.4245}\ \Delta w_{1t} + \underset{(0.0458)}{0.2073}\ \Delta_3 w_{1t-1} - \underset{(0.054)}{0.5011}\ \Delta_3 w_{2t-1} - \underset{(0.00553)}{0.03664}\ \widehat{ec}_{2t-1}$ $+\underset{(0.0853)}{0.07038}\ \Delta_4 p_t + \underset{(0.0792)}{0.1422}\ \Delta_4 p_{t-1} - \underset{(0.00476)}{0.01784}\ \Delta u_{t-2}$
Sector 3 (public sector)
$\Delta w_{3t} = \underset{(0.0787)}{0.1659}\ \Delta w_{1t} + \underset{(0.117)}{0.2506}\ \Delta w_{2t} - \underset{(0.0467)}{0.08951}\ \Delta w_{3t-1} - \underset{(0.0502)}{0.3811}\ \Delta w_{3t-2}$ $-\underset{(0.0454)}{0.2221}\ \Delta w_{3t-3} - \underset{(0.0311)}{0.1857}\ \widehat{ec}_{3t-1} + \underset{(0.0322)}{0.1578}\ \Delta_4 p_t$
CPI-equation:
$\Delta p_t = -\underset{(0.0807)}{0.1779}\ \Delta p_{t-1} - \underset{(0.0766)}{0.02421}\ \Delta p_{t-2} + \underset{(0.0727)}{0.129}\ \Delta p_{t-3} + \underset{(0.0736)}{0.2619}\ \Delta p_{t-4}$ $-\underset{(0.0166)}{0.1047}\ \widehat{ec}_{4t-1}$
mis-specifications tests† and encompassing (the VAR) test‡:
$F_{AR} = 1.1666[0.1748]$ $\chi^2_{NORM} = 13.001[0.1118]$ $\chi^2_{ENC-VAR} = 83.025[0.0653]$
Notes
Sample 1980(1)-2011(4). Estimation is by FIML. Deterministic terms are omitted Standard errors are in parentheses below the parameter estimates. $\widehat{ec}_{i_{t-1}}$ $(i = 1, 2, 3, 4)$ are explained in the text. †System versions of 1-5 order autocorrelation test and normality test, see Doornik and Hendry (2013) †The likelihood-ratio test of the over-identifying restriction, see Doornik and Hendry (2013) The numbers in [ ] are p-values of these tests

The quarterly rate of change in the manufacturing wage rate is seen to depend negatively on the change in the payroll tax rate. This implies that tax increases are rolled back to the wage earners. Since $\widehat{ec}_{1t}$ includes $w_{1t} + \tau_t - q_{1t} - z_{1t}$, the wage level increases one-for-one with a reduction in the payroll tax rate. This means that the hourly wage cost is independent of the payroll tax rate in the long run. This result has been found in earlier empirical studies, also on aggregated data. The negative coefficient of the three-quarter lagged change in the wage rate ($\Delta_3 w_{1t-1}$) is also known from earlier studies, see for example Nymoen (1989a). It reflects the fact that most of the adjustment of the wage level takes place in the same quarter

each year (the second). The estimated effect of an expected rise in the annual inflation rate ($\Delta_4 p_t$) of 1 percentage point is as high as 0.3462. The effect is strongly significant with an implied t-value of 8.08. The implication for the wage norm is discussed in the main text.

The two last variables in the manufacturing wage equation are the unemployment rate (u_t), in log form, and net labour immigration (IM_t) in per cent. The unemployment effect is well known from previous studies and represents a moderately convex wage curve, see e.g. Blanchflower and Oswald (1994), Hoel and Nymoen (1988) and Nymoen and Rødseth (1998). The effect of IM_t could not have been estimated on samples that do not include the massive inflow of labour immigrants since 2005.

Wage growth in the private service sector is strongly influenced by both the (expected) wage norm (Δw_{1t}) and by lagged manufacturing wage increases ($\Delta_3 w_{1t-1}$). The coefficient of the lagged relative wage with respect to manufacturing ($\hat{ec}_{2t-1}$) is also sizeable and significantly different from zero. Taken together, this is strong evidence in favour of wage-following behaviour. The finding that the annual CPI inflation rate enters at one lag, and with a lower estimated parameter than in manufacturing, also indicates that wage adjustments in the private service sector are mainly anchored by the wage norm. That said, we also estimate significant effects of unemployment and immigration in the wage equation for the private service sector, with about same sized coefficients, as in the manufacturing sector wage equation.

The public sector wage equation shows contemporaneous effects of quarterly wage increases in both sector 1 and sector 2, but with more weight on the wage in the private sector. The wage relative to sector 2 is significant, which is consistent with the wage settlements in the public sector being last in the chain. We pick up a similar effect of the expected rise in the cost of living as in sector 2, which therefore emerges as a systemic feature of Norwegian wage formation. The same is true for the effect of the rate of unemployment. A permanent increase in the rate of unemployment lowers the relative wages in the sector, which is consistent with Figure 3.

The structural wage-setting model implies a restricted reduced form maximised likelihood that can be compared to the maximised likelihood of the unrestricted system, or VAR. Since we have a set of over-identifying restrictions, the statistical validity of the ordered wage-fixing system can be tested with respect to the unordered wage-setting system (the VAR) by a likelihood ratio test. This is the $\chi^2_{ENC-VAR}$ reported at the end of Table 9, and the interpretation of the p-value of 0.26 is that the ordered system represents no significant loss of explanatory power relative to the unrestricted and unordered VAR model. The unordered purely statistical system is encompassed by the structural model, see Hendry and Mizon (1993) and (Hendry 1995, Chapter 14). Bårdsen et al. (2005, Ch. 3) presents an application to Norwegian wage-price dynamics of this econometric approach.

We next turn to the detailed results for the dynamic specification of the 1980(1)-2000(4) dataset, cf. Table 10 before inflation targeting was formally introduced in 2001 and while labour immigration to Norway was at a low level. As noted in the main text, it is reasonable to include the unemployment rate and the immigration rate as unrestricted stationary variables on this sample, so that the equilibrium correction terms (denoted $\widehat{ecs}_{it}$ $i = 1, 2, 3$) become the wage share for sector 1, and the two relative wage rates for sectors 2 and 3.

Table 10: The dynamic specification of the leader-followership model in Table 8. Sample 1980(1) - 2000(4).

SECTOR 1 (MANUFACTURING)
$\Delta w_{1t} = -\underset{(0.0345)}{0.1026}\ \Delta\tau_1 - \underset{(0.0498)}{0.27}\ \Delta_3 w_{1t-1} - \underset{(0.0196)}{0.04604}\ \widehat{ecs}_{1t-1} + \underset{(0.0622)}{0.3152}\ \Delta_4 p_t - \underset{(0.00488)}{0.01443}\ u_{t-1} - \underset{(0.0826)}{0.00003}\ IM_t$
SECTOR 2 (PRIVATE SERVICES)
$\Delta w_{2t} = \underset{(0.09)}{0.1531}\ \Delta w_{1t} + \underset{(0.0675)}{0.2177}\ \Delta_3 w_{1t-1} - \underset{(0.0831)}{0.5042}\ \Delta_3 w_{2t-1} - \underset{(0.0949)}{0.1139}\ \widehat{ecs}_{2t-1} + \underset{(0.134)}{0.1079}\ \Delta_4 p_1 + \underset{(0.124)}{0.09343}\ \Delta_4 p_{t-1} - \underset{(0.00396)}{0.01391}\ u_{t-1} - \underset{(0.00569)}{0.01472}\ \Delta u_{t-2} - \underset{(0.0602)}{0.0314}\ IM_{t-4}$
SECTOR 3 (PUBLIC SECTOR)
$\Delta w_{3t} = \underset{(0.14)}{0.4935}\ \Delta w_{1t} + \underset{(0.14)}{0.4935}\ \Delta w_{2t} - \underset{(0.0582)}{0.03097}\ \Delta w_{3t-1} - \underset{(0.0581)}{0.3398}\ \Delta w_{3t-2} - \underset{(0.0484)}{0.1608}\ \Delta w_{3t-3} - \underset{(0.0823)}{0.2443}\ \widehat{ecs}_{3t-1} + \underset{(0.0425)}{0.1585}\ \Delta_4 p_t - \underset{(0.00296)}{0.01027}\ u_{t-1}$
CPI-EQUATION:
$\Delta p_t = -\underset{(0.102)}{0.01002}\ \Delta p_{t-1} - \underset{(0.104)}{0.04616}\ \Delta p_{t-2} - \underset{(0.0884)}{0.2634}\ \Delta p_{t-3} + \underset{(0.0863)}{0.3841}\ \Delta p_{t-4} - \underset{(0.0184)}{0.07642}\ \widehat{ec}_{4t-1}$
MISSPECIFICATIONS TESTS† AND VAR ENCOMPASSING TEST‡:
$F_{AR} = 1.1046[0.2890]$ $\chi^2_{NORM} = 14.251[0.0755]$ $\chi^2_{ENC-VAR} = 143.08[0.0000]$
Notes
Sample 1980(1)-2000(4). Estimation is by FIML. Deterministic terms are omitted Standard errors are in parentheses below the parameter estimates. $\widehat{ecs}_{it}$ $(i = 1, 2, 3)$ and $\widehat{ec}_{4t}$ are explained in the main text.: †System versions of 1-5 order autocorrelation and normality tests, see Doornik and Hendry (2013) ‡The likelihood-ratio test of over-identifying restrictions, see Doornik and Hendry (2013) The numbers in [ ] are p-values of these tests

The similarity of the results for the short-run wage leadership-followership sample compared to the full sample results is discussed in the main text. The implied long-run wage equations from the short-sample results become:

(22) $$w_1 = \tau_t - x_{1t} - 0.31u - 0IM$$

(23) $$w_2 = w_1 - 0.12u - 0.28IM$$

(24) $$w_3 = w_2 - 0.04u$$

which are comparable to, e.g., Panel 4 in Table 4. The main difference is that the estimated effect of immigration on the wage-leading sector is zero on the short-sample, which

is reasonable since "normal" labour immigration to Norway probably has little impact on the bargaining power of the trade unions in the wage-leading sector. In the private wage-following sector, there is an estimated effect, which is also reasonable since there is more direct market regulation and weaker unions in this sector.

References

: n.d.

Aukrust, O.: 1977, Inflation in the open economy. A Norwegian model, *in* L. B. Klein and W. S. Sâlant (eds), *World Wide Inflation. Theory and Recent Experience*, Brookings, Washington D.C.

Bårdsen, G., Eitrheim, Ø., Jansen, E. S. and Nymoen, R.: 2005, *The Econometrics of Macroeconomic Modelling*, Oxford University Press, Oxford.

Bårdsen, G., Jansen, E. S. and Nymoen, R.: 2003, Econometric inflation targeting, *Econometrics Journal* **6**, 429—460.

Blanchard, O. and Galí, J.: 2010, Labor markets and monetary policy: A new keynesian model with unemployment., *American Economic Journal: Macroeconomics* **2**(2), 130.

Blanchard, O. J.: 1987, The wage price spiral, *Quarterly Journal of Economics* **101**, 543–565.

Blanchflower, D. G. and Oswald, A. J.: 1994, *The Wage Curve*, The MIT Press, Cambridge, Massachusetts.

Bowitz, E. and Cappelen, Å.: 2001, Modelling incomes polcies: Some norwegian experiences 1973-1993, *Economic Modelling* **18**, 349–379.

Bratsberg, B. and Raaum, O.: 2012, Immigration and wages: Evidence from construction, *The Economic Journal* **122**, 1177–1205.

Calmfors, L. and Seim, A. L.: 2013, Pattern bargaining and wage leadership in a small open economy, *Scandinavian Journal of Economics* **115**(1), 109–140. DOI: 10.1111/j.1467-9442.2012.01731.x.

Calvo, G.: 1983, Staggered prices in a utility maximizing framework, *Journal of Monetary Economics* **12**, 383–98.

Cuikerman, A. and Lippi, F.: 1999, Central bank independence, centralization of wage bargaining inflation and unemployment: Theory and some evidence, *European Economic Review* (43).

Doornik, J.: 2003, Asymptotic tables for cointegration tests based on the gamma-distribution approach. Nuffield College, University of Oxford.

Doornik, J. A.: 1998, Approximations to the asymptotic distributions of cointegration tests, *Journal of economic surverys* **12**(5), 573—593.

Doornik, J. A. and Hendry, D. F.: 2013, *Modelling Dynamic Systems PcGive 14. Volume 2*, Timberlake Consultants, London.

ECB: 2012, Euro area labour markets and the crisis, *Occasional Paper Series 138*, European Central Bank.

Edgren, G., Faxén, K. O. and Odher, C. E.: 1969, Wages, growth and distribution of income, *Swedish Journal of Economics* **71**, 133–60.

Forder, J.: 2014, *Macroeconomics and the Phillips Curve Myth*, Oxford Studies in the History of Economics, Oxford University Pres, Oxford.

Forslund, A., Gottfries, N. and Westermark, A.: 2008, Prices, productivity and wage bargaining in open economies, *Scandinavian Journal of Economics* **110**((1)), 169—195.

Gertler, M. and Trigari, A.: 2009, Unemployment fluctuations with staggered nash wage bargaining, *Journal of Political Economy* **177**(1), 3886.

Gjelsvik, M., Nymoen, R. and Sparrman, V.: 2013, Lnnsrelasjoner i kvarts og modag, *Technical Report 25*, Statistics Norway.

Hall, R. E.: 2005, Employment fluctuations with equilibrium wage stickiness., *American Economic Review* **95**(1), 50–65.

Harbo, I., Johansen, S., Nielsen, B. and Rahbek, A.: 1998, Asymptotic inference on cointegrating rank in partial system, *Journal of Business & Economic Statistics* **16**, 388–399.

Hendry, D. F.: 1995, *Dynamic Econometrics*, Oxford University Press, Oxford.

Hendry, D. F. and Doornik, J. A.: 2014, *Empirical Model Discovery and Theory Evaluation. Automatic Selection Methods in Econometrics*, Arne Ryde Memorial Lectures, MIT Press, Cambridge,Massachusetts.London, England.

Hendry, D. F. and Mizon, G. E.: 1993, Evaluating dynamic econometric models by encompassing the VAR, *in* P. C. B. Phillips (ed.), *Models, Methods and Applications of Econometrics*, Basil Blackwell, Oxford, pp. 272–300.

Hicks, J.: 1955, Economic foundations of wages policy, *Econimic Journal* **65**(259), 388–404.

Hoel, M. and Nymoen, R.: 1988, Wage formation in Norwegian manufacturing. An empirical application of a theoretical bargaining model, *European Economic Review* **32**, 977–997.

Holden, S.: 2005, Monetary regimes and the co-ordination of wage setting., *European Economic Review* (49), 833–843.

Johansen, K.: 1995a, Norwegian wage curves, *Oxford Bulletin of Economics and Statistics* **57**, 229—247.

Johansen, S.: 1995b, *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*, Oxford University Press, Oxford.

Johansen, S.: 2002, A small sample correction for the test of cointegrating rank in the vector autoregressive model, *Econometrica* **70**(5), 19291961. doi:10.1111/1468-0262.00358.

Krogh, T.: n.d., Real wage rigidity and the unemployment volatility puzzle in small open economies. Statistics Norway and Norges Bank.

Kwiatowski, D., Phillips, P., Schmidt, P. and Shin, Y.: 1992, Testing the null hypothesis of stationary against the alternative of a unit root, *Journal of Econometrics* (54), 159–178.

Layard, R., Nickell, S. and Jackman, R.: 1991, *Unemployment*, Oxford University Press, Oxford.

Marginson, P.: 2014, Coordinated bargaining in europe: From incremental corrosion to frontal assault, *European Journal of Industrial Relations* . Online version published 15 April 2014.

Meade, J.: 1982, *Wage-Fixing, Stagflation Volume 1*, George Allen and Unwin.

Nickell, S. J. and Andrews, M.: 1983, Unions, real-wages and employment in Britain 1951-79, *Oxford Economic Papers (Supplement)* **35**, 183–206.

Norges, B.: 2002, Inflation report, *Technical Report 3*, Norges Bank.

Nymoen, R.: 1989a, Modelling wages in the small open economy: An error-correction model of Norwegian manufacturing wages, *Oxford Bulletin of Economics and Statistics* **51**, 239–258.

Nymoen, R.: 1989b, Wages and the length of the working day. a empirical test based on Norwegian quarterly manufacturing data, *Scandinavian Journal of Economics* **91**, 599–612.

Nymoen, R.: 1991, A small linear model of wage- and price-inflation in the Norwegian economy, *Journal of Applied Econometrics* **6**, 255–269.

Nymoen, R. and Rødseth, A.: 1998, Modelling manufacturing wages in the nordic countries. Unpublished manuscript, University of Oslo.

OECD: 2012, Oecd economic outlook norway, *Technical report*, OECD.

OECD: 2014, *Sharing the pain equally? Wageadjustments during the crisis and recovery*, OECD Publishing, chapter 2. url = "/content/chapter/empl$_o utlook - 2014 - 5 - en$", .

Samuelson, P. A.: 1951, Economic theory and wages, *in* D. Wright (ed.), *The Impact of the Union*, Harcourt Brace & Company, pp. 312–43.

Solow, R. M.: 1990, *The Labour Market as a Social Institution*, Basil Blackwell, Oxford.

Soskice, D. and Iversen, T.: 2000, The nonneutrality of monetary policy with large price or wage setters, *Quarterly Journal of Economics* (1), 265–284.

Usher, D.: 2012, Bargaining unexplained, *Public Choice* **151**(1/2), 23–41.

Visser, J.: 2013, Wage bargaining institutions-from crisis to crisis, *European Economy. Economic Papers 488*, European Commision.

Have inflation targeting and EU labour immigration changed the system of wage formation in Norway?*

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Abstract

Collective agreements have played a central role in the system of wage formation in Norway for more than fifty years. Although the degree of coordination achieved has been variable, pattern wage bargaining has been a mainstay of the system. We investigate the degree of invariance in wage formation in Norway with respect to two recent structural changes: the transition towards inflation targeting in monetary policy and an unprecedented surge in labour supply due to higher immigration rates. We report empirical results that support the view that a semi-permanent high immigration may affect wages negatively in a significant way. However we do not find evidence that the stability of the arbitration system, and in particular the wage-bargaining pattern, has been changed by labour immigration or by inflation targeting monetary policy. An explanation of why we do not find evidence of structural changing effects of the transition of monetary policy, can be found in the fact that the wage arbitration system itself has syncronized the inflation expectations of the social partners. In that analysis, inflation targeting became a new layer of nominal stabilization, on top of the existing one.

Keywords: *Inflation modelling, pattern wage bargaining, inflation targeting, dynamic econometrics, cointegration, small open economy*

JEL classification: *C52, E24, E31, E37, J31.*

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1 Introduction

In Norway, the post-war system for wage formation was originally designed to balance the need for relative inflation control with the attainment of full employment and a balanced current account, cf. Aukrust (1977). The system has survived both the international stagflation of the 1970s and the Norwegian housing and banking crisis in 1989-90. It was based on an agreement between the government and the wage setters. The government committed to a fixed exchange rate policy, while the wage setters committed to collective wage bargaining. The system of wage fixing in Norway unfolds along both the horizontal and vertical dimension, implying that wage setting is based both on bargaining between evenly matched labour unions and firms' organizations, and bargaining at a high level. Later, similar systems were adopted in, for example, Sweden, see Edgren et al. (1969) and the Netherlands. A presentation of the Norwegian model from an econometric perspective is found in Bårdsen et al. (2005, Ch. 3).

The system for wage formation created a link between the wage norm for the overall economy, which was determined by the sector exposed to international competition, and profitability and productivity growth in the same sector. For several decades the system has linked wage increases in the other sectors of the economy (known as wage-followers) to wage formation in the exposed sector (the wage-leader). Down the years, it has been accepted, even supported, by governments of different party colours. Probably a recognition of the idea that this version of pattern bargaining represents an operational way of creating a certain degree of coordination in wage setting.

One threat to this system is the secular decline in the traditional manufacturing industries, which in Norway has been accelerated by the extraction of oil and gas and the use of the resulting 'oil-money'. Nevertheless, several commentators have noted the stability of the Norwegian norm-based system. For example by Visser (2013, p. 59-60), and also by OECD economists:

> ... the so called traditional ('non-oil') sector has been diminishing in importance. Despite this, it is important in wage setting. Rather than wages being determined by the relative bargaining strength of different sectors, the general wage level is set by the social partners first considering the wage increases that the traditional sector can "afford". OECD (2012, p. 15)

Nevertheless, the functioning of the wage formation system has been challenged by two important events in the last decade: the transition towards inflation targeting formally introduced in 2001 and the EU enlargement in 2004, which led to a large increase in the labour supply.

The change in monetary policy, from exchange rate stabilization to inflation targeting, can potentially change wage formation and the degree of coordination among wage setters, see Cuikerman and Lippi (1999), Soskice and Iversen (2000) and Holden (2005). In Norway, the introduction of inflation targeting in 2001 could potentially have a direct impact on wage formation, to the extent that it came into conflict with the 'existing order' based on manufacturing sector wage leadership. This potential for conflict was brought into focus by the central bank's *Inflation Report* from 2002. That report stated that the new monetary

policy required a reversal of the causality flow in wage formation.[1] The bank's worry was that the wage leader model could pull inflation in one direction (up for example, because productivity growth is higher there than in the service sectors), at the same time as monetary policy analysis suggested a reduction of aggregate demand relative to the bank's estimate of the Norwegian economy's potential output. This issue can also be related to theoretical research, showing that the pattern of wage bargaining under inflation targeting may be of marginal importance, e.g. Calmfors and Seim (2013). Hence, the hypothesis emerged that new monetary policy would force a change of roles in the pattern wage bargaining, where the wage-followers may become wage-leaders and vice versa.

With regards to immigration, research based on microeconomic data has found evidence that immigration has reduced wage earnings in certain groups in the labour market, see Bratsberg and Raaum (2012). The hypothesis is that, unless the immigrants emulate incumbent workers as regards trade union membership, increased immigration may reduce trade unions' bargaining power. The result might be that wage growth in sectors with high immigration rates deviates from the general wage growth.

We investigate empirically the degree of invariance in wage formation with respect to these structural changes. First, we investigate whether the value of labour productivity remains the main determinant of the wage trend in the exposed sector. Second, we test whether the wage norm created by the exposed sector still has a defining role in Norwegian wage formation. Third, we investigate whether inflation expectations have become more important after the introduction of inflation targeting. Finally, we include a labour immigration variable in our information set and identify its role in our structural model of wage formation.

We use quarterly data from 1980 to 2011. This means that about 1/3 of the sample comes from the era with a monetary policy regime that targeted inflation, and that about 1/4 of the sample comes from the recent period of increasing and high labour immigration. Hence, if either the monetary policy regime shift or the new immigration flow has affected wage formation or the social order with any force, the evidence should be in this data set.

The paper is organised as follows. In Section 2, we outline the theoretical framework of coordination in Norway and define the twin concepts of wage-leader and wage-follower sectors in wage formation. Based on the theoretical and statistical properties of the time series in Section 3, we formulate an empirical model, test the wage-leader/-follower properties, and carry out analysis of the model in Section 4. In Section 5, we conclude that the econometric results suggest that the Norwegian wage model has preserved its main properties in spite of the recent challenges to the system.

[1] The following quotation from Inflation Report number 3/2002, page 28 summarizes the hypothesis: *With an inflation target for monetary policy and a floating exchange rate, it is the inflation target, not wage growth abroad, that determines the level of growth which is consistent with stable profitability in the business sector over time. Inflation in Norway will over time be determined by the inflation target that the Government has set for monetary policy and not inflation abroad. Exchange rate developments are determined by inflation differentials between Norway and other countries.*

2 Wage formation

The twin concepts of wage-leader and wage-followers are important in the Norwegian system of wage formation. Wage-leader refers to the sector (or sectors) where the bargaining outcome defines the wage norm that forms the basis for bargaining in the other sectors of the economy (wage-followers). Throughout the whole post-war period, the settlements in the manufacturing sector have in practice defined the wage norm.[2]

We also define the manufacturing industry as the wage-leading sector but we test whether that role has been preserved so far in the new millennium. For the wage-followers, we have chosen to distinguish between the private service sector (including construction, mainly because of the importance of engineering consultancy in this sector) and the public sector. But also because negotiations in the public sector start after the wage settlements in private sector. Hence, in the following, we number the sectors 1 (wage-leader), 2 (wage-follower, private) and 3 (wage-follower, public).

2.1 The Wage-leader

The formulation of a theory of a sustainable wage norm requires an assessment of not only self-interest among workers and firms, but also of compromise. As pointed out by Usher (2012), 'compromise is then not just another way of talking about self-interest, and social, political and institutional forces are not merely cover-ups for imprecisely modelled individuals rational actions, they are among the fundamental determinants of decisions. In this view, even a full analysis of rational behaviour leads to an indeterminacy of wages, and other considerations had to be introduced to resolve it.

The recognition among economists that there is an indeterminacy in the economic theory of wages goes back to the 1950s, see Forder (2014, Ch. 1.4) who cites Samuelson (1951, p. 312) and Hicks (1955, p. 390) and other leading theorists. The economic theory of supply and demand could set some limits to what wages can be set, but within those limits closer determination requires that other relationships are introduced. The indeterminacy of wages from theory also characterizes the now standard Diamond-Mortensen-Pissarides (DMP) search and matching model. In the DMP model, the wage is usually determined in a Nash bargaining game. But is the wage logically equal to the Nash solution given the assumptions of the DMP model? As Hall (2005) pointed out, any wage in the bargaining set is in principle consistent with private efficiency on the part of both the firm and the worker. In that sense, the equilibrium wage rate is only set-identified. He then went on to analyze a solution where the real wage is fixed, which however is only one possibility of what in the DMP-literature is referred to as wage 'stickiness'. [3]

[2]Because of its importance, the system of wage-formation is a recurrent theme in public (and at times academic) debate. For example, the "traditional" definition of wage-leader has been challenged several times over the last decades. Those wanting to "change the content" of the wage-leading sector has emphasized both reduced legitimacy of the norm, as the importance of the sector for the total wage bill has become smaller, and reduced relevance (since Norwegian manufacturing may have become integrated in the super profitable oil and gas extraction). Among the arguments that are used in favour of no-change is that a manufacturing wage-norm aids coordination.

[3]Following Hall (2005), several papers have incorporated rigid wage setting in search models. For instance, Gertler and Trigari (2009) present a DMP model where the frequency of wage bargaining is constrained by

At the same time as we find it challenging to determine wages theoretically, we also observe that actual wage bargains are struck year after year, and that they are rationalized by considerations of profits, actual and required (to attract investments), cost of living and relative wages (fairness). These observed regularities give reason to believe that wage formation can be subject to econometric treatment. This is also what have motivated much of the econometric literature recently surveyed by Forder (2014). A development that goes back to the first half of the 1960s in Norway (although it was documented in English only later, by Aukrust (1977)) was that the wage norm became defined as a long-run trend, and also that short-run deviations had to be tolerated as 'part of the system'. But how should the trend be defined? Should it be a combination of the many things that might legitimate a secular trend trend? Or should the number of norm determined factors be limited? In Norway, a view that won support on both sides for the arbitration system, was that it was important to 'keep it simple' and relate the wage norm to the value of labour productivity in the exposed sector of the economy.

This idea of a wage norm can be treated econometrically using cointegration methods and equilibrium correction modelling. One of the main implications that we follow up is that, measured on a logarithmic scale, the wage level in the exposed sector, should be cointegrated with the log of the level of product prices and average labour productivity. Long-run price homogeneity together with a unit long-run elasticity on productivity lead to the implication of a stationary wage-share which may serve as an operationalization of an equilibrium wage-share. In the wider interpretation, the equilibrium income distribution needs not be completely constant but can depend on intermittent shifts, or trends, in bargaining power (possibly proxied by the unemployment rate) and the support to the idea about necessary compromise and coordination.

We base our econometric model on the assumption that both product price (q_1) and average labour productivity (z_1) are random-walk processes with drift. It follows that the time series for w_{1t} (the logarithm of the hourly wage in the wage-leader sector) contains both a deterministic and a stochastic (random walk) trend.[4]

Formally, the value of average labour productivity can be split into two components: q_{1t} (the log of the price in Norwegian currency) and z_{1t} (the log of average productivity in fixed prices), and these exogenous processes can be expressed as two random-walks with drift:

$$q_{1t} = q_{10} + q_{1t-1} + v_{q_1 t} \tag{1}$$

$$z_{1t} = z_{10} + z_{1t-1} + v_{a_1 t} \tag{2}$$

(1) and (2) imply that q_1 and z_1 are integrated of order 1, denoted $I(1)$. Positive deterministic trends in prices and productivity require $q_{10} > 0$ and $z_{10} > 0$.

Note that we could write equation (1) in terms of a price that is denoted in foreign currency, q^f_{1t}, and the exchange rate xr_t (log of kroner per currency unit) and specify time

Calvo (1983)) style lottery, leading to sticky wages. Blanchard and Galí (2010) combine a reduced form of search model with real wage rigidity with a New Keynesian model to study how this impacts monetary policy. Krogh (n.d.) generalizes the Hall-approach to a small open economy model where there is a non-trivial distinction between the consumer real wage and the producer real wage.

[4]In this section, we abstract from payroll tax rates in all three sectors. A representative payroll tax is included in the empirical model.

series models for q^f_{1t} and xr_t. To be consistent with $q_{10} > 0$ in equation (1), one or both of these time series models must contain a positive drift parameter. We assume that the price level variable q^f_{1t} has a positive drift, and for simplicity, we also assume that it is a constant parameter. The specification of the process for the nominal exchange rate is more complicated. For the period with a fixed exchange rate regime, it can possibly be specified without drift, but with intermittent structural breaks in order to represent devaluations. After the switch to the present floating exchange rate regime, a random walk, possibly with drift, is probably reasonable for the nominal exchange rate.

Hence, also across exchange rate regimes, equation (1) with $q_{10} > 0$ is a reasonable first approximation of the evolution of the price level in the wage-leading sector. As noted, there may be structural breaks in the drift parameter, and the foreign exchange rate regime-shift may also mean v_{q_1t} is not white-noise. The main point, however, is that the unit root property is robust across regimes and, in particular, that there are no explosive roots that logically must be included in the model of wage-price formation during the period with a floating exchange rate.

As noted, we assume that q_{1t} and z_{1t} jointly determine the deterministic trend in w_{1t}. In the simplest case, w_{1t} is cointegrated with the sum $q_{1t} + z_{1t}$, meaning that the logarithm of the wage share is I(0) with a constant mean s_1. This relationship can be written as:

$$w_{1t} - q_{1t} - z_{1t} = s_1 + e_{1t} \tag{3}$$

$$e_{1t} \sim I(0), \tag{4}$$

The sum of the productivity trend and the foreign price trend plays an important role in the framework, since it traces out a central tendency or a long-run sustainable scope for wage growth. This wage norm seems to correspond well to the concept of a wage corridor (Aukrust (1977)) for wage determination in the industries that are most exposed to foreign competition.

Cointegration implies equilibrium correction. Therefore, with reference to equations (1) and (2), e_{1t-1} should have significant predictive power for wage growth, Δw_{1t} ($\equiv w_{1t} - w_{1t-1}$):

$$\begin{array}{lcl} a) & (w_1 - q_1 - z_1)_{t-1} & \rightarrow \Delta w_{1t} \\ b) & (w_2 - w_1)_{t-1} & \nrightarrow \Delta w_{1t} \\ c) & (w_3 - w_1)_{t-1} & \nrightarrow \Delta w_{1t} \end{array} \tag{5}$$

The economic interpretation is that collective wage bargaining, through its focus on the distribution of value added between labour and capital, implies that an equilibrium wage share (the parameter s_1) is maintained over time, cf. Forslund et al. (2008). (5b) and (5c) capture the idea that if the manufacturing sector is wage leading, it is implied that Δw_{1t} cannot be influenced by the lagged relative wage to the two other sectors.

As mentioned above, the equilibrium wage share is likely to depend on several underlying factors, e.g. related to production technology, product market conditions and bargaining power. It is only when these factors are constant that s_1 is likely to be a stable parameter. More generally, we can therefore add a time subscript to s_1 in equation (3) and write s_{1t} as a function of the variables that we include as its determinants in the econometric model, namely the log of the unemployment rate, u_t, and the immigration rate, IM_t:

$$s_{1t} = s_{10} + \beta_{11} u_t + \beta_{12} IM_t.$$

Hence, the extended long-run wage equation for sector 1 that we use in the following becomes

$$w_{1t} - q_{1t} - z_{1t} = s_{10} + \beta_{11} u_t + \beta_{12} IM_t + e_{1t} \tag{6}$$

If we abstract from IM, (7) is a standard linearised wage bargaining model, with full weight on producer prices and no weight on consumer prices, see e.g. Nickell and Andrews (1983) and Hoel and Nymoen (1988). The signs of the parameters, β_{11} and β_{12}, are expected to be negative (or zero). A higher unemployment rate can reduce the bargaining power of the unions. A marked change in the immigration rate may also affect union bargaining power and coordination negatively, unless a large proportion of the immigrants choose to become union members.

2.2 Wage-followers

If the manufacturing wage level represents the wage norm in the pattern of wage bargaining, and if $w_{1t} \sim I(1)$, w_{2t} must also be $I(1)$ and cointegrated with w_{1t}:

$$w_{2t} - w_{1t} = s_2 + e_{2t} \ , \tag{7}$$

$$e_{2t} \sim I(0), \tag{8}$$

where s_2 is the equilibrium relative wage. Note that equation (7) and (8) can be maintained by adjustments to wages in both sector 1 and sector 2. In order to define sector 2 as a wage-follower with sector 1 as a wage-leader we require:

$$\begin{array}{lccc} a) & \Delta w_{1t} & \overset{\longrightarrow}{\nleftarrow} & \Delta w_{2t} \\ b) & (w_{2t-1} - w_{1t-1}) & \longrightarrow & \Delta w_{2t} \end{array} \tag{9}$$

Requirement (9a) implies that the contemporaneous relationship between the wage growth rates in the two sectors is recursive. We can implement and test (9a) as a restriction on a simultaneous equations model of wage setting. (9b) requires that the stationarity of the relative wage is due to equilibrium correction in sector 2.

In the Norwegian system of wage fixing, bargaining in the public sector starts after negotiations in the private sector, and it is therefore natural to assume that w_{3t} equilibrium corrects with respect to w_{2t}, i.e.:

$$w_{3t} - w_{2t} = s_3 + e_{3t} \tag{10}$$

$$e_{3t} \sim I(0). \tag{11}$$

The required recursive structure is the same as above, but between sector 2 and sector 3:

$$\begin{array}{lccc} a) & \Delta w_{2t} & \overset{\longrightarrow}{\nleftarrow} & \Delta w_{3t} \\ b) & (w_{3t-1} - w_{2t-1}) & \longrightarrow & \Delta w_{3t} \end{array} \tag{12}$$

Clearly, for the government sector to be follower of sector 2, Δw_2 in (9) should not respond to $(w_3 - w_2)_{t-1}$. In both sectors, we include the possibility of potential wage shifting in the empirical model. Hence, s_2 in equation (7) and s_3 in equation (10) can include effects from the unemployment rate and from the immigration flow in the same way as in equation (6).

2.3 Wage-price inflation

The wage-price spiral is well known from the literature, cf. Blanchard (1987), Meade (1982) and Layard et al. (1991). In the wake of the financial and jobs crisis, the dynamics of wage and price setting have come to be regarded as central to rebalancing the euro area after the crisis, cf. ECB (2012) and OECD (2014).

Wage increases to compensate for expected increases in the cost of living are regularly demanded in the collective wage bargaining context, also in the wage-leading manufacturing sector. In Norway, a government-supported body, the Technical Calculation Committee for Wage Settlements (TCC), has the anchoring of CPI inflation forecasts as one of its tasks. The committee consists of representatives of both employer and employee organisations, and the consolidation of expectations takes place before the annual rounds of wage fixing starts.

Since CPI inflation, in turn, depends on growth in wage costs (in addition to the price of imported consumer goods), it is clear that a realistic system of equations for Δw_{1t}, Δw_{2t} and Δw_{3t} will not be completely recursive: in the short run, wage adjustments in the manufacturing sector also depend on CPI inflation expectations.

Nevertheless, as long as the wage norm that defines e_{1t} does not include the level of CPI, the inclusion of the rate of inflation in the system of wage formation represents a realistic modification of the otherwise recursive structure.

2.4 Potential reversal of bargaining pattern

As noted in the introduction, it is possible that inflation targeting caused an inversion of the causal ordering in wage formation. Specifically, that the relative wage $w_1 - w_2$ becomes a predictor of Δw_1, rather than of Δw_2. According to this hypothesis, neither (5b) nor (9b) will hold empirically on the latter sample period when monetary policy has changed. The hypotheses of reversed causality can be tested within the cointegrated VAR .The VAR is estimated in section 4.

3 Temporal properties of the time series variables

In this section, we first examine the assumption that wages, prices and productivity are non-stationary. Second, we examine whether the residuals of the equilibrium correction relationships are stationary. Stationary residuals between integrated variables imply a cointegrating relationship between the variables in the long-term solution.

Time series for wages, productivity and producer price developments are taken from the National Accounts. The unemployment rate is from the labour force survey, while consumer prices, immigration and population figures are collected from the official pages of Statistics Norway. Variables measured in natural logarithms are denoted by lower case Latin letters. Appendix A contains a detailed variable description.

3.1 The manufacturing sector

Wages, product prices and productivity in the manufacturing sector have increased over time, see Figure 1. The statistical properties of the variables are tested using two unit

root tests, the Augmented Dickey Fuller test (ADF) and a test by Kwiatowski et al. (1992) referred to as KPSS henceforth. The null of the ADF test is a unit root, and a rejection of the null implies that the time series is stationary. The null of the KPSS, on the other hand, is stationarity. Hence a rejection implies that the time series is integrated. Both tests are used to shed light on the time series properties.

The KPSS test rejects that w_1 is stationary, see Table 1. The ADF test statistic (including a constant, trend and seasonal dummies) is just above the critical value, which suggests the opposite result. The KPSS test result is supported by the testing of stationarity of the first difference of the same variable. The ADF test statistic of w_1 in difference (including a constant and seasonal dummies) is clearly above the critical value. Results from both ADF and KPSS tests indicate that the time series for product prices, q_1 is integrated. The ADF test statistic for productivity, z_1, does not reject I(1), while the KPSS test indicates that stationarity cannot be rejected. While the opposite is true for consumer prices, p_t, the ADF test rejects I(1), while the KPSS test strongly rejects stationarity. However, it is logically inconsistent to regard CPI inflation as I(1), while interpreting wages and import prices as stationary. In sum, we treat the variables as integrated of order 1 in the econometric modelling.

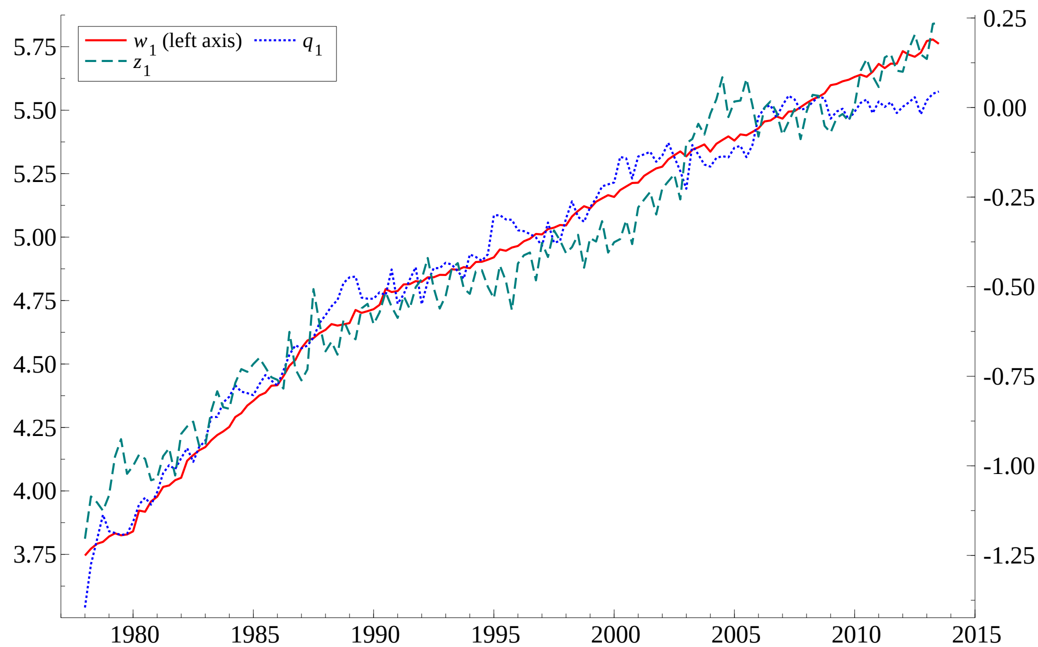

Figure 1: Components of the manufacturing wage share.

Figure 2 shows the unemployment rate, u_t, and the immigration (flow) rate , IM_t. Both variables are percentages (see Appendix A). The unemployment rate and the immigration rate both reject stationarity according to the KPSS test, see Table 1. The graph shows both the first period of departure from post-war full employment in the first half of the 1980s, and the return to near full employment in the years following deregulation of the financial markets (among other things). Next, unemployment increased again during the

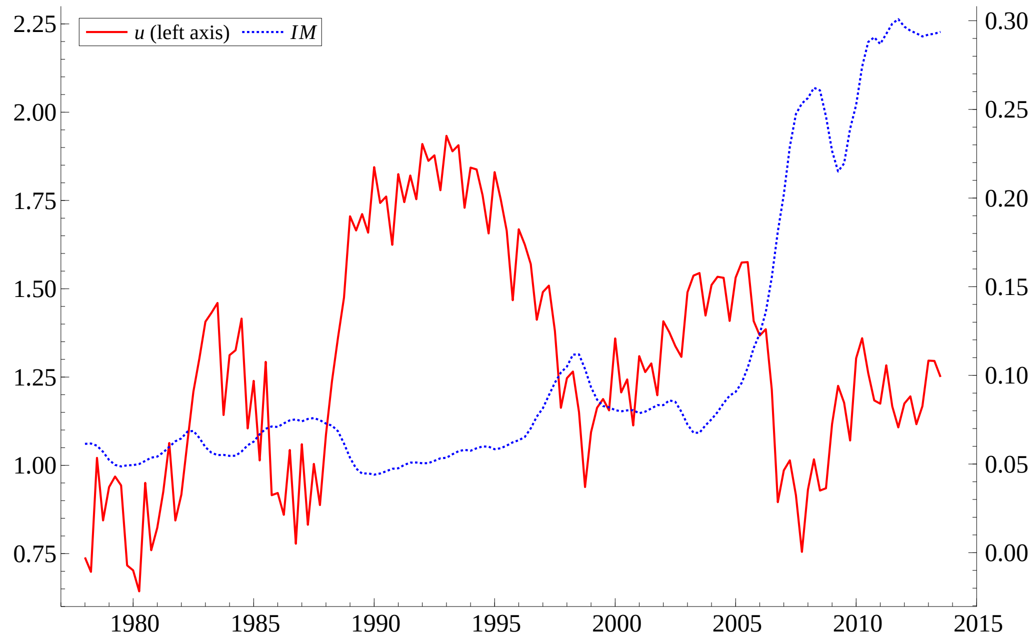

Figure 2: Plot of the logs of the unemployment rate and the immigration flow.

1990s following the fall in oil prices, a restrictive fiscal policy, the collapse of house prices and the ensuing banking crisis. Finally, the small changes in unemployment since 2005 are interesting. The number of employed wage earners in Norway grew by an unprecedented quarter of a million from 2005 to 2009. Naturally, the rate of unemployment fell, but not nearly as much as employment increased. This is due to the high immigration rates. This graph shows labour immigration to Norway from the EU and other developed countries. The migration inflow amounts to 1.6 per cent of the total labour supply in 2011 and is clearly large enough to significantly affect Norwegian labour markets and therefore possibly also wage setting.

We have also investigated the properties of the residuals from a regression of equation (3) and (6), see Table 2. The residuals from regression of equation (3) are labelled e_1, while the residuals from regression of equation (6) are labelled $e_{1,UIM}$. The KPSS test statistic for the wage share in Table 2 is lower and below the critical value if we control for shifts in labour immigration and unemployment. The tests therefore imply that the wage share is stationary when these variables are included. The ADF test does not reject the hypothesis of a unit root, neither for e_1 nor for $e_{1,UIM}$.

3.2 The public sector and private services

We have also investigated the statistical properties in the two other sectors; public and private services. Stationarity is rejected by the KPSS test for the hourly wage rate in both sectors. Figure 3 shows wages in sector 2 relative to sector 1, e_2, together with wages in sector 3 relative to sector 2, e_3. While wages in sector 2 appear to follow wages in sector 1 quite

Table 1: Unit root tests on the variables in levels and first differences. Sample period 1980(1) to 2011(4).[1]

Level				First difference			
Variable	ADF (Lags)	KPSS	Char.[2]	Variable	ADF (Lags)	KPSS	Char.[2]
w_1	-3.714 (1)	0.282	c, t, s	Δw_1	-3.232 (4)	0.878	c, s
z_1	-2.255 (3)	0.076	c, t, s	Δz_1	-10.22 (2)	0.055	c, s
q_1	-2.382 (4)	0.263	c, t, s	Δq_1	-5.550 (3)	0.520	c, s
u	-2.810 (4)	0.205	c, t, s	Δu	-3.841 (3)	0.160	c, s
p	-5.979 (4)	0.310	c, t, s	Δp	-1.842 (3)	1.009	c, s
IM	-0.728 (2)	0.260	c, t, s	ΔIM	-5.352 (1)	0.401	c, s
w_2	-4.303 (4)	0.260	c, t, s	Δw_2	-2.659 (3)	0.818	c, s
w_3	-4.290 (0)	0.192	c, t, s	Δw_3	-9.680 (3)	0.760	c, s

[1]Note, however, that the sample period varies due to the number of lags in the ADF test.

[2]The characteristics are a constant (c), a trend (t) and seasonal dummies (s).

The 5% critical value of the ADF test is -3.45 and -2.88 in level and differences respectively.

The 5% critical value of KPSS test is 0.146 in level and 0.463 in differences.

closely, the wage level in sector 3 was markedly lower than the wage level in sector 2 for a long period. Although this reduction in the relative wage started in late 1980s, it is palusible that it was related to the rise in unemployment in the 1990s and the tight government budgets that characterised the first half of the decade, in particular. This indicates that, in the econometric model, the rate of unemployment can help to explain the development in relative wages. In the last period, wages in sector 3 have increased relative to wages in sector 2, while wages in sector 2 have decreased relative to sector 1. This is the period of high labour immigration.

In Table 2, the residuals from the regression of (7) are labelled e_2, while the residuals from a regression of (7), which includes u and IM, are labelled $e_{2,UI}$. The same regressions are performed for sector 3 as well, and residuals are labelled e_3 and $e_{3,UI}$, respectively. According to the ADF tests in Table 2, e_2, ec_{2UI} and ec_3 are non stationary. Again, however, the KPSS test supports the stationarity of the residuals and indicates that taking account of the possibility that unemployment and immigration can influence the long-run relationships, results in more stationary residuals.

Table 2: Unit root tests on the wage share and the relative wages. Sample period 1980(2) to 2011(4).

Variable	ADF (Lags)	KPSS	Char.[1]
e_1	-2.833 (3)	0.583	c
$e_{1,UIM}$	-2.751 (3)	0.305	c
e_2	-2.439 (3)	0.428	c
$e_{2,UIM}$	-2.232 (3)	0.155	c
e_3	-2.135 (3)	0.328	c
$e_{3,UIM}$	-3.153 (3)	0.170	c

The 5% critical value of the ADF test is -2.88 and 0.463 for the KPSS test
[1]The tests have included a constant (c) in the test.

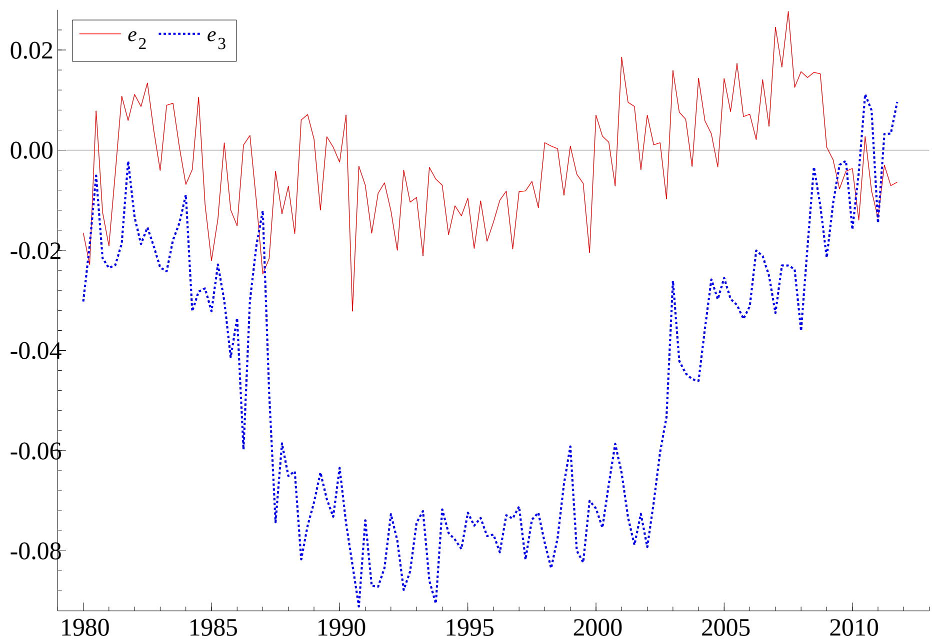

Figure 3: Wages in sector 2 relative to sector 1 (e_2) and wages in sector 3 relative to wages in sector 2 (e_3).

4 Econometric results

The previous section showed that univariate tests give clear support for unit-root non-stationarity, and some initial support for three cointegration relations of the pattern bargaining model. However, cointegration is a system property, and a concise way of testing the long-run theory is to formulate a vector autoregression (VAR) and apply the maximum likelihood method of Johansen (1995b). This step in the analysis is taken in section 4.1.

In section 4.2, we specify a cointegrated equilibrium correction model (ECM) that is consistent with the evidence from the VAR-based test of cointegration. The ECM is used to further test the hypotheses about the dynamics of wage leader- and follower bargaining model (the causal structure between the growth rates). The tests illustrate the relative invariance of the model parameters after the change from a fixed exchange rate to a floating exchange rate and the surge in immigration rates.

4.1 Cointegration

We formulate a fourth order VAR for the three wage levels, manufacturing (w_1), private services (w_2) and public sector (w_3). We include the endogenous variables of the VAR in the vector $\mathbf{Y_{wt}} = (w_{1t}, w_{2t}, w_{3t})'$. The VAR is expressed as:

$$\Delta \mathbf{Y_{wt}} = \sum_{i=1}^{3} \mathbf{\Gamma}_i \Delta \mathbf{Y_{wt-i}} + \mathbf{\Pi} \mathbf{Y_{t-1}} + \mathbf{\Upsilon} \mathbf{Q_t} + \varepsilon_{\mathbf{t}}, \tag{13}$$

where all variables in the cointegration space are defined in the extended data vector: $\mathbf{Y_t} = (\mathbf{Y_{wt}}, x_t, \tau_{1t}, u_t, IM_t)'$. $x = q_1 + z_1$ is the wage-scope variable in the manufacturing sector. The role of the payroll tax rate, τ_{1t}, enables us to distinguish between the wage rate per hour and the wage cost per hour. $\mathbf{\Gamma}_i$ $(i = 1, \ldots, 4)$ and $\mathbf{\Pi}$ are matrices with short-run and long-run parameters for wages, unemployment and immigration.

Formally, τ_{1t}, u_t and IM_t are also treated as I(1), since they are included in the $\mathbf{Y_t}$ vector. However, the tax rate is determined by political decisions and apparent unit root non-stationarity is probably due to policy determined changes in the mean. As the discussion above suggests, while u_t and IM_t are clearly non-stationary, the economically relevant 'model of stationarity' is probably that they are also driven by intermittent structural breaks in their means, rather than being convincingly represented by unit-root non-stationarity.

$\mathbf{Q_t}$ is other non-modelled variables in the VAR, namely the lagged annual CPI inflation rate and dummy variables. For simplicity, they are associated with the single matrix of parameters $\mathbf{\Upsilon}$. Finally, the vector $\varepsilon_{\mathbf{t}}$ contains the VAR residuals for the three wage growth rates.

As noted above, $\mathbf{Q_t}$ in (13) implies that the VAR includes other variables as well. In terms of economic interpretation, the most important variable is the lagged rate of consumer price inflation, which is included to control for the effects of cost of living considerations (as noted above). In the sample period, there have been periods of centralised wage settlements (wage laws), see e.g. Bowitz and Cappelen (2001) and Section 3. We include dummies for these events, and for other quarters with outliers. Finally, we include a constant, three centred seasonal dummies and a deterministic trend. The result is a relatively well-specified VAR. The residuals, $\hat{\varepsilon}_{\mathbf{t}}$, have a tendency to be negatively auto-correlated, indicating a certain over-fitting. However, the test of joint non-normality of $\varepsilon_{\mathbf{t}}$ is insignificant at the 5 % significance level.

$\mathbf{\Pi}$ is often referred to as the long-run matrix, and it has the dimensions 3×7. Let r denote the rank of $\mathbf{\Pi}$. Since x_t, τ_{1t}, u and IM are non-modelled variables, r can be 0, 1, 2 or 3. $r = 0$ means that there is no cointegration, which would imply rejection of the theory we

formulated above. If, on the other hand, $r = 3$, the variables x_t, τ_{1t}, u and IM represents four common trends in the three nominal wage rates. The theory of wage leadership points to the wage-scope variable $x_t = q_{1t} + z_{1t}$ as the most dominant trend in the system. This is because the evolution of the scope variable defines the upward trend in the wage corridor in the manufacturing sector. If wage leadership holds, this trend is transmitted to the two other wage rates.

$\mathbf{\Pi}$ can be written as $\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where $\boldsymbol{\alpha}$ is $3 \times r$ and $\boldsymbol{\beta}$ ($7 \times r$) is the matrix with cointegration parameters. The theory presented in section 2 implies $\tau = 3$ with (3), (7) and (10) as the cointegration relationship.

When conducting the cointegration rank test a trend is included and restricted to lie in the cointegration space (i.e. included in $\mathbf{Y_{t-1}}$) such that critical value exist, see Harbo et al. (1998). Table 3, column *Trace test statistic* presents the test statistic of the cointegration analysis. The critical values of Doornik (2003) for zero, one and two cointegrating relationships are presented in Table 3, column five. The trace test statistic for two or fewer cointegration relationships is below the 10 % critical value for a standard integrated VAR reported in Doornik (2003). This implies that the test rejects two or fewer cointegration relationships at the 10 % significance level. This result supports the hypothesis of three cointegrating relationships. However, these critical values do not take into account exogenous variables. The critical values increase with the number of exogenous variables included, cf. Harbo et al. (1998). The critical values with four exogenous variables are shown in Table 3, last column.

The two first test statistics are well above their respective critical values (adjusted for exogenous variables), which supports that the number of stationary long-run relationships is at least two. The last row shows that we are unable to formally reject two, and accept three, cointegrating vectors. However, the critical values are for asymptotic distributions, and using these critical values may lead to over-rejecting in small samples, see Doornik (1998). The correction factor for critical values without exogenous variables is given in Johansen (2002). Correction factors do not exist for critical values in cointegration analyses with exogenous variables, but the correction factor without exogenous factors implies that the critical values of Harbo et al. (1998) should be interpreted with care. We conclude that cointegration tests, broadly interpreted, provide formal support for three cointegrating relationships ($r = 3$), but that equilibrium correction may be relatively weak.

Table 4, Panel 1, shows the maximised likelihood value, *Log L*, after we have omitted the deterministic trend from the three long-run relationships. The trend was included for the purpose of making correct inferences about the rank, but once the rank has been fixed to three, its significance can be tested by a standard Likelihood-Ratio (LR) test with a Chi-square distribution. In this case, the restriction is rejected at the 1 % level. However, keeping a trend is without meaning economically, since the wage share and relative wages will then drift deterministically. A closer inspection of the results also reveal that the trend is sensitive to if the first three years in public sector is included. There was a substantial wage stop in public sector which ended in the beginning of the 80s and wages increased quite rapidly in the following years. Hence, if the first three years are excluded the tests imply that the trend should be removed. We therefore conclude that the trend should be removed from the cointegration space. In Table 4, we use the result of a rank equal to 3, but use a VAR without trend when testing the identification of the cointegrating relationships.

Table 3: Tests of cointegration rank. Sample 1982(1)-2011(4).

Eigenvalue (λ_i)	H_0	H_1	*Trace* *Test statistic*	10% *Critical value Doornik* *No exogenous*	*With exogenous*
0.37	$r = 0$	$r >= 1$	116.5	39.7	60.5
0.23	$r <= 1$	$r >= 2$	51.3	23.3	37.7
0.17	$r <= 2$	$r >= 3$	14.3	10.7	18.5

Endogenous variables: wages in manufacturing (w_1), private services (w_2) and public sector (w_3).
Restricted variables: wage scope ($x = q_1 + z_1$) in manufacturing,
unemployment (u), immigration (IM) and a trend (t).
Unrestriced variables: payroll tax-rate, τ_1, constant, seasonal dummies and some specific dummies which capture important events in the Norwegian wage setting, see e.g. Bowitz and Cappelen (2001) and Section 4.

Removing the deterministic trend does not contribute to identification. In Panel 1, we therefore show the results after we have imposed identifying restrictions that are consistent with the wage leadership theory. Specifically, the payroll tax rate is only included in the cointegrating relationship, which is normalised on w_1. This separates the second and third relationships from the first. The exclusion of x from the second and third relationships provides more identifying information for these two relationships. In order to separate the second from the third relationship, we exclude w_1 and IM from the third, and w_3 from the second relationship. Finally, in order to identify the first relationship, both w_2 and w_3 are excluded from the relationship that is normalised on w_1. The identified long-run relationships are reported in Panel 1 of the table, together with the likelihood value and the Chi-square distributed test of the seven over-identifying restrictions that represent the wage leadership view of Norwegian wage formation. The restrictions are statistically acceptable at the 5 % level.

Table 4: Testing steady-state hypotheses

Unrestricted system for wages

$w_1 = \beta_{w2,1}w_2 + \beta_{w3,1}w_3 + \beta_{\tau,1}\tau + \beta_{x,1}x + \beta_{u,1}u + \beta_{IM,1}IM$

$w_2 = \beta_{w1,2}w_1 + \beta_{w3,2}w_3 + \beta_{\tau,3}\tau + \beta_{x,2}x + \beta_{u,2}u + \beta_{IM,2}IM$

$w_3 = \beta_{w1,3}w_1 + \beta_{w2,3}w_2 + \beta_{\tau,3}\tau + \beta_{x,3}x + \beta_{u,3}u + \beta_{IM,3}IM$

Unrestricted $Log\ L = 1365.79768$

Panel 1: Identified long-run relationships

$w_1 = \underset{(.)}{-1}\,\tau + \underset{(.)}{1}\,x \underset{(0.09)}{-0.38}\,u \underset{(0.42)}{-1.72}\,IM$

$w_2 = \underset{(.)}{1}\,w_1 \underset{(0.07)}{-0.40}\,u \underset{(0.20)}{-0.51}\,IM$

$w_3 = \underset{(.)}{1}\,w_2 \underset{(0.014)}{-0.07}\,u$

$Log\ L = 1359.27709$, $\chi^2(7) = 13.04[0.07]$

Panel 2: Manufacturing wage leadership, $\alpha_{w1,2} = \alpha_{w1,3} = 0$

$w_1 = \underset{(.)}{-1}\,\tau + \underset{(.)}{1}\,x \underset{(0.09)}{-0.41}\,u \underset{(0.43)}{-1.76}\,IM$

$w_2 = \underset{(.)}{1}\,w_1 \underset{(0.07)}{-0.40}\,u \underset{(0.21)}{-0.57}\,IM$

$w_3 = \underset{(.)}{1}\,w_2 \underset{(0.015)}{-0.08}\,u$

$Log\ L = 1359.25748$, $\chi^2(9) = 13.08[0.16]$

Additional restrictions: $\chi^2(2) = 0.04(0.98)$

Panel 3: No response of Δw_2 to $(w_{2t-1} - w_{3t-1})$, $\alpha_{w2,3} = 0$

$w_1 = \underset{(.)}{-1}\,\tau + \underset{(.)}{1}\,x \underset{(0.09)}{-0.41}\,u \underset{(0.43)}{-1.76}\,IM$

$w_2 = \underset{(.)}{1}\,w_1 \underset{(0.07)}{-0.42}\,u \underset{(0.19)}{-0.55}\,IM$

$w_3 = \underset{(.)}{1}\,w_2 \underset{(0.016)}{-0.08}\,u$

$Log\ L = 1359.18919$, $\chi^2(10) = 13.22[0.21]$

Additional restrictions: $\chi^2(1) = 0.14[0.71]$

Panel 4: No direct response of Δw_3 to manufacturing norm, $\alpha_{w3,1} = 0$

$w_1 = \underset{(.)}{-1}\,\tau + \underset{(.)}{1}\,x \underset{(0.12)}{-0.58}\,u \underset{(0.53)}{-2.45}\,IM$

$w_2 = \underset{(.)}{1}\,w_1 \underset{(0.03)}{-0.17}\,u \underset{(0.15)}{-0.18}\,IM$

$w_3 = \underset{(.)}{1}\,w_2 \underset{(0.01)}{-0.04}\,u$

$Log\ L = 1358.39232$, $\chi^2(11) = 14.81[0.19]$

Additional restrictions: $\chi^2(1) = 1.59[0.21]$

Conditional on the identified long-run relationships, we can begin to test for the minimum restrictions that imply no feed-back from w_2 and w_3 to w_1. The results are reported in Panel 2 of the table, using $\alpha_{i,j}$ to denote the element in row i, column j in the matrix with loadings parameters. The increase in the test of the over-identifying restrictions is small. Therefore it is not surprising that two zero restrictions that distinguish Panel 2 from Panel 1, are statistically acceptable (the incremental test yields $\chi^2(2) = 0.04[0.98]$). We note that the estimated cointegration parameters also change very little from Panel 1 to Panel 2.

We conclude that the tests in Panel 1 and 2 provide relatively clear support for the view that the wage-leading role of the manufacturing sector has been in operation during our sample period. The result is different from those estimated by Nymoen (1991), who was unable to impose the wage-leading restrictions when estimating on a sample that started in 1969(1) and ended in 1987(4). As a result, he specified a preferred model that allowed for wage-wage effects. The different samples might be relevant here, since the 1970s in particular were characterized by super high wage inflation in certain years. Some of this may have been driven by low coordination and wage-wage inflation.[5]

Returning to Table 4, Panel 3 and 4 test additional restrictions implied by pattern wage bargaining. In panel 3, we add the restriction that the deviations from the second and third cointegration relationships are not corrected by manufacturing wages. This result shows that the increase in the test of over-identifying restrictions is small, and the incremental test does not reject additional restrictions. The cointegration parameters are also very similar when we compare Panel 2 and Panel 3. Finally, in Panel 4, the restriction that wages in sector 3 does not equilibrium correct with respect to the first cointegrating relationship is imposed. Although this restriction is acceptable, we note that the parameter estimates change notably. We return to the question of a possible contemporaneous effect of w_1 on w_3 in the structural model below.

The point estimates in the best identified models, e.g. in Panel 3, deserve comment. In the manufacturing sector (Sector 1), the elasticity of wages ($\beta_{u,1}$) with respect to unemployment is higher in absolute value than the estimates in, e.g., Nymoen (1989) and Johansen (1995a), who used data from the 1980s.

This may be due to the larger variation of u_t in our sample, which includes the period after the Norwegian banking crisis in the 1990s. For Sector 2, the estimated unemployment elasticity is practically the same ($\hat{\beta}_{u,2} = -0.4$). We note, however, that this implies that the estimated long-run wage responsiveness for the private wage-following sector is twice as large as for the wage-leading sector. This is realistic, since this sector is characterized by weaker union organisation and lower capacity for multi-employer agreements than in the manufacturing sector. In the public sector, the estimated parameter of unemployment ($\beta_{u,3}$) is much lower (in absolute value) than in the two other sectors, which is also reasonable given that changes in overall unemployment have little direct relevance for the two bargaining parties.

Since labour immigration is a new phenomenon, there are no studies that can serve as direct references for discussing our results for the parameters $\beta_{IM,1}$ and $\beta_{IM,2}$ in Panel 3.

[5]There are other differences as well, both in the data definitions and in the econometric methodology. Nymoen did not identify the cointegrating relationships separately from the short-run dynamics, as we will do. This may have led to poorer identification of his long-run relationships.

However, one important study that use micro data reports evidence that immigration has causally reduced wage earnings in certain sectors of the Norwegian economy, see Bratsberg and Raaum (2012). Compared to our own earlier research using single equation estimation methods, the estimated value of $\hat{\beta}_{IM,1} = -1.7$ is very large, see Gjelsvik et al. (2013). Note, however, that the estimated standard error indicates a broad 95 % confidence interval. The single equation estimation just mentioned provides point estimates of somewhat less than -1 for the wage-leading sector, and this value is within the confidence interval with good margin.

Based on the above interpretation of the evidence we have specified the following equilibrium correction variables for use in a structural econometric model:

$$\widehat{ec}_{1t} = w_{1t} + \tau_t - x_t + 0.38u_t + 1.1IM_t \tag{14}$$

$$\widehat{ec}_{2t} = w_{2t} - w_{1t} + 0.40u_t + 0.51IM_t \tag{15}$$

$$\widehat{ec}_{3t} = w_{3t} - w_{2t} + 0.07u_t \tag{16}$$

The only significant difference from Panel 1, is that we have adjusted the point estimate of the coefficient of IM in (14). The estimation results of the dynamic structural model will confirm that this downsizing of the direct effect of immigration on the manufacturing wage is data-acceptable.

4.2 A dynamic simultaneous equations econometric model

In this section, we formulate and estimate a dynamic model model that is consistent with the cointegration analysis, and which is identified econometrically. At this stage in the analysis we include CPI inflation into the model in an explicit way, which was suggested to be important to wage formation in Section 2.

In the VAR above, the lagged four-quarter change in the log of the official consumer price index, denoted $\Delta_4 p_{t-1}$ was included among the variables in $\mathbf{Q_t}$ in (13). In the same way as for the lagged wage growth rates, we can think of lagged price inflation as a variable in a reduced form obtained from a simultaneous equations model, SEM, that includes price inflation among the endogenous variables. In order to model price growth jointly with wage growth, we need a separate long-run relationship for the price level. We do not carry out a cointegration analysis for the price level in this paper. Instead, we use the following equilibrium correction term

$$\widehat{ec}_{4t} = p_t - 0.6(w_{1t} + \tau_t - z_{1t}) - 0.4pi_i \tag{17}$$

where the elasticities for unit wage cost and import price, pi, are taken from the econometric results in Bårdsen et al. (2003).

The SEM model can formally be expressed as:

$$\mathbf{B_0 Y_{wpt}} = \sum_{i=1}^{3} \mathbf{B}_i \mathbf{Y_{wpt-i}} + \mathbf{AEC_{t-1}} + \Upsilon \mathbf{Q_t} + \epsilon_\mathbf{t}, \tag{18}$$

where $\mathbf{Y_{wpt}} = (\Delta w_{1t},\ \Delta w_{2t},\ \Delta w_{3t}, \Delta_4 p_t)'$ contains changes in wages and the consumer price index, $\mathbf{EC}$ contains the long run relationships in (14)-(17) and $\mathbf{Q_t}$ is redefined to exclude

consumer prices. $\mathbf{A}$ is a diagonal matrix and all the elements of the diagonal in $\mathbf{B_0}$ is equal to one.

The last four columns in table 5 show coefficient estimates ($\hat{\mathbf{A}}$) for an identified four equation SEM, where a diagonal matrix with adjustment coefficients for the four ec-terms is sufficient for identification on the rank condition.[6] The maximised log likelihood reported in the last row serves as the unrestricted likelihood value that we use to test the wage-leadership-followership restrictions in Table 6.[7] The two test statistics are (F_{AR}), for autocorrelation of order between 1 and 5, and non-normality (χ^2_{NORM}), calculated from the SEM residuals. They are system versions of autocorrelation and normality tests, see Doornik and Hendry (2013), and are reported with asymptotic p-values in square brackets.

Table 5: Wage-price SEM. FIML estimates of the contemporaneous coefficient matrix $\hat{\mathbf{B}}_\mathbf{0}$ (first four columns with estimates) and adjustment coefficients $\hat{\mathbf{A}}$ (last four columns). Standard errors in brackets below the estimates. Sample period 1980(1) to 2011(4).

	Δw_{1t}	Δw_{2t}	Δw_{3t}	$\Delta_4 p_t$	$\widehat{ec}_{1t-1}$	$\widehat{ec}_{2t-1}$	$\widehat{ec}_{3t-1}$	$\widehat{ec}_{4t-1}$
Δw_{1t}	-1	-0.05 (0.15)	-0.04 (0.10)	0.37 (0.06)	-0.04 (0.01)	0	0	0
Δw_{2t}	0.42 (0.06)	-1	0.03 (0.06)	0.05 (0.09)	0	-0.04 (0.006)	0	0
Δw_{3t}	0.17 (0.08)	0.24 (0.12)	-1	0.16 (0.03)	0	0	-0.19 (0.03)	0
$\Delta_4 p_t$	0.02 (0.06)	-0.12 (0.12)	0.09 (0.09)	-1	0	0	0	-0.10 (0.02)

Log L= 1892.0816, $F_{AR} = 1.18[0.16]$ and $\chi^2_{NORM} = 12.6[0.12]$.

Table 5 shows that the diagonal elements of the matrix with adjustments coefficients for the ec-terms are statistically significant, although the numerical values are relatively small for the first two. This is consistent with the results in Table 3 and Table 4. The results for the contemporaneous parameters ($\hat{\mathbf{B}}_\mathbf{0}$) also show a clear pattern. In the row for Δw_{1t}, the two coefficients of Δw_{2t} and Δw_{3t} are negative ("wrong sign"), but they are statistically insignificant from zero values (judged by the standard errors). Conversely, Δw_{1t} has a sizeable estimated coefficient in the rows for Δw_{2t} and Δw_{3t}, which supports the hypothesis that the manufacturing sector is wage-leading, with private service production and the public sector as wage-followers.

The first row in Table 6 shows that three versions of the null hypothesis that manufacturing is *not* wage-leading, are formally rejected at the 5 % level. In the second line of tests, the second entry supports that the private service sector is wage-leading in relation to the public sector. The third row of tests shows that the coefficient of Δw_{3t} can be restricted to

[6] At this stage, there are no restrictions on the correlation matrix of the contemporaneous disturbances, which means that the rank condition is necessary and sufficient.

[7] Note that this likelihood is not comparable to the corresponding in Table 4, since the model in Table 5 includes CPI inflation as an endogenous variable

Table 6: Likelihood-ratio tests of wage-leader/follower restrictions on the model in Table 5

Restrictions:	Sec 1 $\nrightarrow$ Sec 2	Sec 1 $\nrightarrow$ Sec 3	Sec1 $\nrightarrow$ Sec 2 and 3
	$\chi^2(1) = 54.5^{**}$	$\chi^2(1) = 4.6^{*}$	$\chi^2(2) = 54.9^{**}$
Restrictions:	Sec 2 $\nrightarrow$ Sec 1	Sec 2 $\nrightarrow$ Sec 3	Sec 2 $\nrightarrow$ Sec 1 and 3
	$\chi^2(1) = 0.11$	$\chi^2(1) = 4.4^{*}$	$\chi^2(2) = 4.7$
Restrictions:	Sec 3 $\nrightarrow$ Sec 1	Sec 3 $\nrightarrow$ Sec 2	Sec 3 $\nrightarrow$ Sec 1 and 2
	$\chi^2(1) = 0.19$	$\chi^2(1) = 0.20$	$\chi^2(2) = 0.42$

* and **denotes significance at the 5% and 1 % levels.

zero in both the manufacturing sector wage equation and in the private service sector wage equation without any significant drop in the likelihood value.

We next turn to the role of price inflation as measured by $\Delta_4 p_t$. This variable is significant with a large estimated coefficient in the manufacturing wage equation. The coefficient of inflation in the equation for Δw_{2t} is close to zero, but can note that there is a significant effect of the lagged inflation rate in this equation, see Appendix B with detailed estimation results.

The last row of Table 5, with the results for the inflation equation $\Delta_4 p_t$, shows that there is little support for within-quarter effects of wage changes on inflation.[8] However, since $\widehat{ec}_{4t}$ includes wage costs $(w_{1t}+ \ln(1+\tau_t))$, we nevertheless have a closed feed-back loop between wages and CPI.

Table 7 shows the results of a restricted estimation where we have imposed all of the six exclusion restrictions discussed above. Compared to Table 5, the retained coefficients change very little.

Table 7: Wage-price SEM with wage leader-follower restrictions imposed. FIML estimates of $\hat{\mathbf{B}}_\mathbf{0}$ and $\hat{\mathbf{A}}$ in equation (18) with standard errors in brackets below the estimates. Sample period 1980(1) to 2011(4).

	Δw_{1t}	Δw_{2t}	Δw_{3t}	$\Delta_4 p_t$	$\widehat{ec}_{1t-1}$	$\widehat{ec}_{2t-1}$	$\widehat{ec}_{3t-1}$	$\widehat{ec}_{4t-1}$
Δw_{1t}	-1	0	0	0.35 (0.06)	-0.04 (0.01)	0	0	0
Δw_{2t}	0.42 (0.06)	-1	0	0.06 (0.09)	0	-0.04 (0.01)	0	0
Δw_{3t}	0.17 (0.08)	0.25 (0.12)	-1	0.16 (0.03)	0	0	-0.19 (0.03)	0
$\Delta_4 p_t$	0	0	0	-1	0	0	0	-0.10 (0.02)

$Log\ L$ =1890.98642, $F_{AR} = 1.17[0.17]$ and $\chi^2_{NORM} = 13.0[0.11]$
LR test of wage-leader following restrictions: $\chi^2(6) = 2.19[0.14]$

[8]The estimated SEM includes three unrestricted lags in the quarterly inflation rate in the fourth row. This ensures that there are no restrictions implied by normalisation on $\Delta_4 p_t$ in the fourth row. This representation is convenient for modelling the effects of inflation on wage dynamics.

It is interesting that Table 5 and Table 7 show that inflation expectations[9] have a significant effect on wage formation in the manufacturing sector, and therefore on the wage norm of the system. Is this evidence of inflation targeting in the sense that the wage increases in the manufacturing sector are led by the inflation forecast of the Central Bank rather than by the evolution of the wage scope? In order to investigate this possibility, as well as the invariance of the system with respect to the monetary policy regime shift more generally, we turn to the estimation results on a sample that ends before the structural break in the market for foreign exchange and in interest rate determination.

4.3 Stability and invariance

In order to address the the question of stability of the wage formation structure, we have estimated the model on a sample that ends in 2000(4). This sample ends before the labour immigration from Europe started, and it is unlikely that the immigration effects on wages that we estimated above are invariant to the shortening of the sample.

Consequently, it does not make sense to define the three wage equilibrium correction terms in the same way as in (14)-(16). Instead, we remove IM_t and u_t from the equilibrium correction variables, which for the short sample are defined as:

$$\widehat{ecs}_{1t} = w_{1t} + \tau_t - x_{1t} \tag{19}$$

$$\widehat{ecs}_{2t} = w_{2t} - w_{1t} \tag{20}$$

$$\widehat{ecs}_{3t} = w_{3t} - w_{2t} \tag{21}$$

implying the same long-run relationships as in the full sample. The estimated system is therefore specified by including immigration and unemployment as exogenous I(0)-variables (but subject to breaks) in the system. For the CPI level, we use (17) as the equilibrium correction term also in the short sample.

The results of the estimation with the wage-leadership-follower restrictions imposed are shown in Table 8. The estimation results are very similar to the full sample estimates in Table 7, although the joint test of the six restrictions is marginally significant at the 5 % level. However, and as previously discussed, the test statistic is not corrected for the smaller sample size. The details show that this is mainly due to an estimated negative coefficient of Δw_{2t} in the Δw_{1t} equation, i.e. the same sign problem that we noted for the full sample results, but more pronounced on the short sample. Appendix B contains Table 10 with the detailed dynamic specification for the short sample.

[9] Inflation in period t is included as an endogenous variable in the model estimated by FIML. Therefore, the inflation variable in the manufacturing sector wage equation for example, can be interpreted as an expectations variable.

Table 8: Pre inflation targeting wage-price SEM with wage-leader-follower restrictions imposed. FIML estimates with standard errors in round brackets below the estimates. Sample period 1980(1) to 2000(4).

	Δw_{1t}	Δw_{2t}	Δw_{3t}	$\Delta_4 p_t$	$\widehat{ecs}_{1t-1}$	$\widehat{ecs}_{2t-1}$	$\widehat{ecs}_{3t-1}$	$\widehat{ec}_{4t-1}$
Δw_{1t}	-1	0	0	0.32 (0.06)	-0.05 (0.02)	0	0	0
Δw_{2t}	0.45 (0.06)	-1	0	0.11 (0.13)	0	-0.11 (0.09)	0	0
Δw_{3t}	0.15 (0.09)	0.49 (0.14)	-1	0.16 (0.04)	0	0	-0.24 (0.08)	0
$\Delta_4 p_t$	0	0	0	-1	0	0	0	-0.08 (0.02)

$Log\ L = 1288.14658$, $F_{AR} = 1.10[0.29]$ and $\chi^2_{NORM} = 14.3[0.08]$
LR test of wage-leader following restrictions: $\chi^2(6) = 12.39[0.05]$

The results show in particular that in manufacturing the estimated coefficient of $\Delta_4 p_t$ is 0.32, with a t-value of 5.1. This is very close to the full sample results. Taken together, the two estimates show that the effect of inflation expectations on the wage norm was in place before inflation targeting was introduced. This is not surprising since CPI expectations have been an important part of the wage setting system for decades. Coordination of inflation expectations among representatives of the employer and employee confederations has been one of the main purposes of the TCC institution long before inflation targeting was introduced.

Another way of illustrating the stability of the estimated structure is by forecasting. Figure 4 shows dynamic forecasts for the period 2001(1) to 2011(4). As can be expected, there are some examples of forecast failures (actuals outside the forecast uncertainty fans), in particular for inflation. However, a major structural break in the 40-quarter forecast period would have resulted in much clearer forecast.

Finally, we illustrate relative parameter stability in the dynamic multipliers of the structural model. In Figure 5, the impulse responses to a shock in the inflation equation (an inflation shock) are shown. There are two graphs in each panel, corresponding to the short sample estimation and the full sample estimation. In all panels, the dynamic multipliers are very similar, illustrating that the results of this policy analysis do not depend on the ten years of inflation targeting. The sign of the multipliers changes from positive to negative. The interpretation is that, because the long-run growth rate in the manufacturing sector depends on producer price growth and productivity growth, the short-run influence of a shock to inflation will mean overshooting in manufacturing sector wages. Because of the wage-leader role of manufacturing, the overshooting spills over to wages in the two other sectors.

Figure 6 shows the impulse response parameters of a shock to the immigration rate (IM). These impulse responses are not invariant. As noted above, this is not surprising given that the surge in immigration came after 2000 (cf. Figure 2). Specifically, it is the effects of increased immigration on wage formation in the manufacturing sector that are underestimated on the sample that ends in 2000(4). Since wage growth in the wage-leading sector is important for the overall nominal path of the Norwegian economy, this spills over

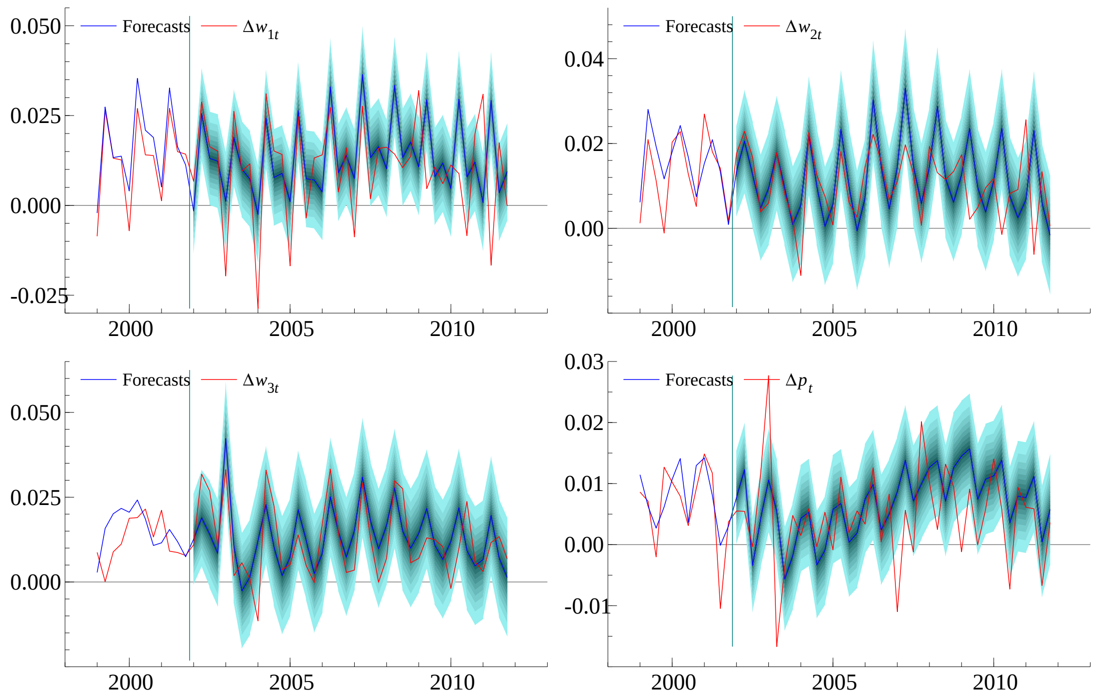

Figure 4: Dynamic forecasts with a 95 per cent forecast uncertainty fans. Forecasts are in blue, actuals in red. The estimation period is from 1980(1) to 2000(4).

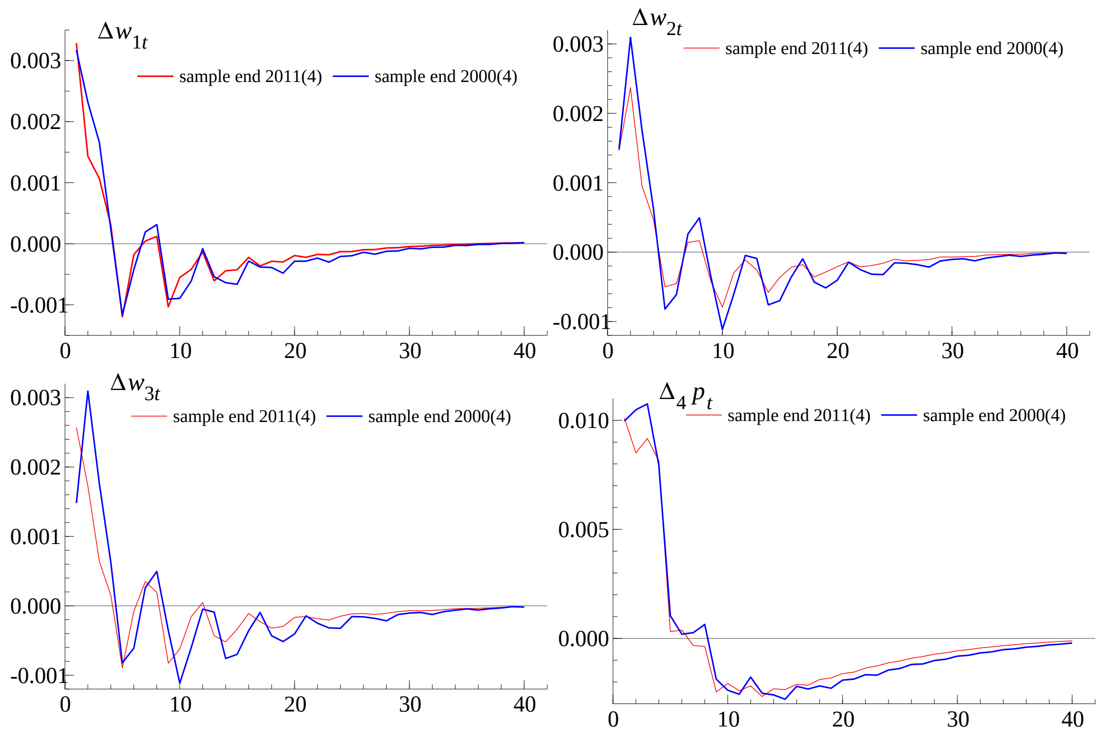

Figure 5: Impulse responses of a shock to the rate of quarterly inflation of 0.01 per cent.

to the inflation impulse responses in particular, which are severely underestimated on the sample that ends in 2000(4).

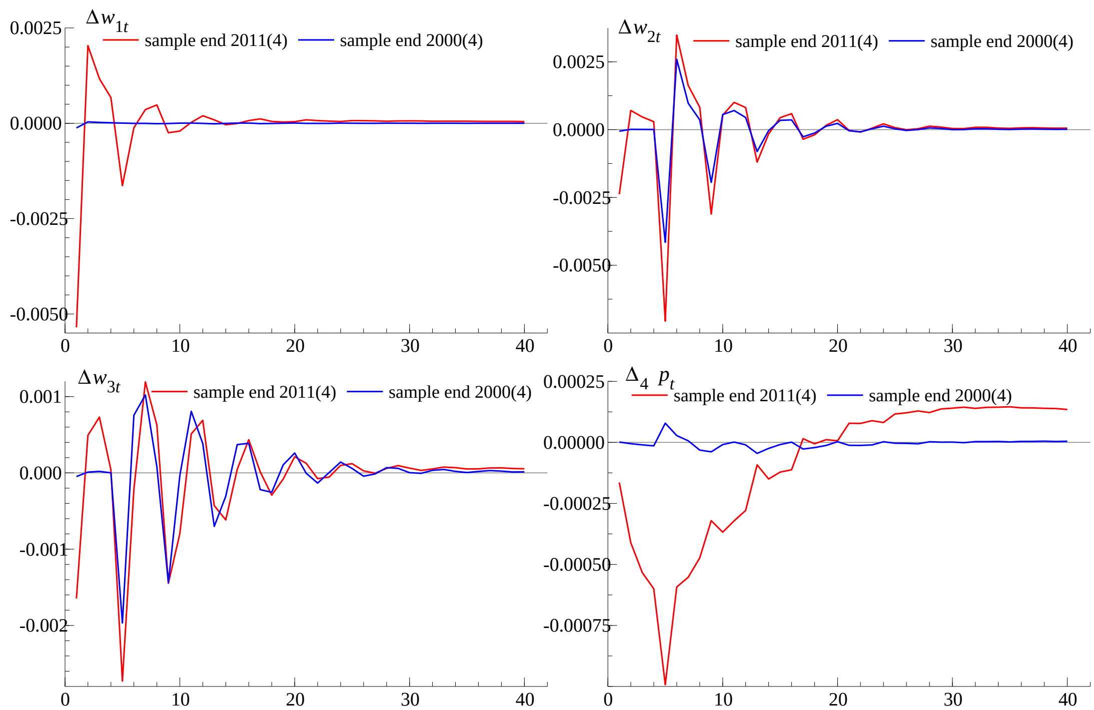

Figure 6: Impulse responses of a shock to the immigration rate by 0.10 percentage points.

5 Conclusion

In this paper we have modelled wages in three sectors of the Norwegian economy; manufacturing, private services and public sector. The model represents the typical pattern of wage bargaining in Norway, where manufacturing sector negotiates first and where the outcome is considered to be the wage norm for the other sectors of the economy. The logic of this pattern is simple, if the wage growth in the exposed sector becomes a wage norm for wage setting in the other sectors, the functional income distribution will be relative both in the wage-leading and wage-following sectors. The wage norm is defined in a system of arbitration that involves self-interest among both workers and firms, but, fundamentally also compromise. As noted in the introduction, the analysis of the system of wage formation as a 'social institution' applies quite, cf Solow (1990). generally. What is perhaps more special about Norway is the continued importance of collective wage agreements, and the roles that the confederate labour market organizations hold the tripartite interaction between government and social partners.[10].

Over the last 10-15 years, there have been two specific challenges to the Norwegian system of wage formation. First, the change to a floating exchange rate and, as a consequence, inflation targeting monetary policy. Second, the rise in labour immigration from outside Scandinavia. Both changes could potentially affect the bargaining power or the gain of coordinating wage formation, and hence change the recursive structure of wage formation in Norway.

The analytical framework that we specify, allows us to test the historical pattern of wage bargaining with a model with no pattern imposed. The model also allows the level of net immigration flow to affect wages separately.

The empirical results give the following answers to the questions posed in the introduction. The cointegration analysis show that there is a stable long term relationship between wage levels in the three sectors, with the manufacturing sector as the wage leader. Therefore, wage growth in manufacturing still constitute the wage norm of the Norwegian wage formation. The empirical results show that the wage norm in manufacturing depends heavily on the profitability growth in the same sector, which is consistent with the theory of collective wage bargaining, but the long term wage level is lower due to the negative and significant effect of the immigration flow. Immigration has also affected the relative wage between private and public services. This might suggest that immigration has affected bargaining power (a parameter of the system), but so far without fundamentally altering the pattern wage bargaining system. And although both unions and employment confederations have understood the nature and aims of monetary policy, we find no indication that the Norwegian monetary policy in 2001 has changed the system of wage formation or that inflation expectations have become more important.

Immigration effects apart, we find that the econometric model has more or less the same parameter estimates on two different samples: one with the period of inflation targeting, and another that ends before this policy change was formally introduced in 2001. The impulse responses of a shock to inflation also show a high degree of invariance. The recursive structure where the wage settlements in manufacturing represent a norm for wage setting is still an

[10] cf. OECD (2012, p. 15)

important part of the Norwegian wage settlements, despite both the structural break in monetary policy and high immigration rates. Hence, the corrosion of, or the attacks on the system of coordinated bargaining analyzed by Marginson (2014) have not been seen in Norway. Instead, the level of coordination is probably notably higher than it was during the 1980s, as also Visser (2013) concludes in his broad analysis of trends in wage bargaining institutions , see Visser (2013, p. 63)

One of the explanations why inflation targeting had little impact on the structure of wage setting, may be that the system of wage formation already included an important element of inflation forecasting and expectations formation. Formally, the Technical Calculation Committee for Wage Settlements has for decades helped to synchronize the inflation anticipations of unions and by firms. For this reason, the central bank's focus on inflation forecast and expectations when inflation targeting started in 2001, was not as "new" to wage and price setters as some commentators would have it.

There are of course other aspects of monetary policy that affect how difficult or easy it is to reconcile inflation targeting with the Norwegian system of wage formation. It is quite likely that a practice of strict inflation targeting (large weight on the inflation-gap and short policy horizon) would have put stress on the system. However, Norges Bank has presented itself as a super flexible inflation targeter. The indication is that the prospects for both unemployment and GDP growth carried large weights in the interest rate policy decisions already before the financial crisis in 2008. During most of the period of inflation targeting, Norges Bank have adopted a relatively long policy horizon, and has shown no haste in closing inflation gaps. The possibility that monetary policy in the future will become more strict to changes in wage setting remains only a vague prospect.

A Data definitions and sources

As explained in the text, lower case letters refer to the logarithm of the original variables listed below. For example, $u_t = \log(U_t)$ denotes the log of the unemployment rate. Variables in first differences are denoted by Δ. Subscripts denote time period. For example, p_{t-4} refers to the (log of) the price level four periods back.

$W_{it}-$Index for hourly wage in sector i=1,2,3
P_t- Consumer price index
$Q_{1t}-$ Price deflator of gross value added, manufacturing industry
$Z_{1t}-$ Labour productivity, output per hour in manufacturing
PI_t- Price deflator on imports of goods and services
U_t- Unemployment rate, in per cent. Civilian unemployment,
IM_t- Immigration from land group 1 and 2, in percent of the population aged 15-74. Group 1 include EU/EFTA countries, North America, Australia and New Zealand. Group 2 includes Eastern Europe except EU countries.
τ_t- represents the natural logarithm of the payroll tax rate plus one.

All variables are from the database of the macroeconometric model KVARTS, maintained by Statistics Norway.

B Additional estimation results

In this appendix, Table 9 shows the detailed dynamic specification of the model in Table 7, while Table 10 is the counterpart to Table 8 in the main text.

Table 9: The dynamic specification of the leader-followership model in Table 7

SECTOR 1 (MANUFACTURING)
$\Delta w_{1t} = -\underset{(0.0368)}{0.1198}\ \Delta\ln(1+\tau_t) - \underset{(0.0489)}{0.3834}\ \Delta_3 w_{1t-1} - \underset{(0.00797)}{0.04055}\ \widehat{ec}_{1t-1} + \underset{(0.042)}{0.3462}\ \Delta_4 p_t$
SECTOR 2 (PRIVATE SERVICE)
$\Delta w_{2t} = \underset{(0.0558)}{0.4245}\ \Delta w_{1t} + \underset{(0.0458)}{0.2073}\ \Delta_3 w_{1t-1} - \underset{(0.054)}{0.5011}\ \Delta_3 w_{2t-1} - \underset{(0.00553)}{0.03664}\ \widehat{ec}_{2t-1}$ $+ \underset{(0.0853)}{0.07038}\ \Delta_4 p_t + \underset{(0.0792)}{0.1422}\ \Delta_4 p_{t-1} - \underset{(0.00476)}{0.01784}\ \Delta u_{t-2}$
SECTOR 3 (PUBLIC SECTOR)
$\Delta w_{3t} = \underset{(0.0787)}{0.1659}\ \Delta w_{1t} + \underset{(0.117)}{0.2506}\ \Delta w_{2t} - \underset{(0.0467)}{0.08951}\ \Delta w_{3t-1} - \underset{(0.0502)}{0.3811}\ \Delta w_{3t-2}$ $- \underset{(0.0454)}{0.2221}\ \Delta w_{3t-3} - \underset{(0.0311)}{0.1857}\ \widehat{ec}_{3t-1} + \underset{(0.0322)}{0.1578}\ \Delta_4 p_t$
CPI-EQUATION:
$\Delta p_t = -\underset{(0.0807)}{0.1779}\ \Delta p_{t-1} - \underset{(0.0766)}{0.02421}\ \Delta p_{t-2} + \underset{(0.0727)}{0.129}\ \Delta p_{t-3} + \underset{(0.0736)}{0.2619}\ \Delta p_{t-4}$ $- \underset{(0.0166)}{0.1047}\ \widehat{ec}_{4t-1}$
MIS-SPECIFICATIONS TESTS† AND ENCOMPASSING (THE VAR) TEST‡:
$F_{AR} = 1.1666[0.1748]$ $\chi^2_{NORM} = 13.001[0.1118]$ $\chi^2_{ENC-VAR} = 83.025[0.0653]$
Notes
Sample 1980(1)-2011(4). Estimation is by FIML. Deterministic terms are omitted Standard errors are in parentheses below the parameter estimates. $\widehat{ec}_{i_{t-1}}$ $(i = 1, 2, 3, 4)$ are explained in the text. †System versions of 1-5 order autocorrelation test and normality test, see Doornik and Hendry (2013) †The likelihood-ratio test of the over-identifying restriction, see Doornik and Hendry (2013) The numbers in [ ] are p-values of these tests

The quarterly rate of change in the manufacturing wage rate is seen to depend negatively on the change in the payroll tax rate. This implies that tax increases are rolled back to the wage earners. Since $\widehat{ec}_{1t}$ includes $w_{1t}+\tau_t-q_{1t}-z_{1t}$, the wage level increases one-for-one with a reduction in the payroll tax rate. This means that the hourly wage cost is independent of the payroll tax rate in the long run. This result has been found in earlier empirical studies, also on aggregated data. The negative coefficient of the three-quarter lagged change in the wage rate ($\Delta_3 w_{1t-1}$) is also known from earlier studies, see for example Nymoen (1989a). It reflects the fact that most of the adjustment of the wage level takes place in the same quarter

each year (the second). The estimated effect of an expected rise in the annual inflation rate ($\Delta_4 p_t$) of 1 percentage point is as high as 0.3462. The effect is strongly significant with an implied t-value of 8.08. The implication for the wage norm is discussed in the main text.

The two last variables in the manufacturing wage equation are the unemployment rate (u_t), in log form, and net labour immigration (IM_t) in per cent. The unemployment effect is well known from previous studies and represents a moderately convex wage curve, see e.g. Blanchflower and Oswald (1994), Hoel and Nymoen (1988) and Nymoen and Rødseth (1998). The effect of IM_t could not have been estimated on samples that do not include the massive inflow of labour immigrants since 2005.

Wage growth in the private service sector is strongly influenced by both the (expected) wage norm (Δw_{1t}) and by lagged manufacturing wage increases ($\Delta_3 w_{1t-1}$). The coefficient of the lagged relative wage with respect to manufacturing ($\hat{ec}_{2t-1}$) is also sizeable and significantly different from zero. Taken together, this is strong evidence in favour of wage-following behaviour. The finding that the annual CPI inflation rate enters at one lag, and with a lower estimated parameter than in manufacturing, also indicates that wage adjustments in the private service sector are mainly anchored by the wage norm. That said, we also estimate significant effects of unemployment and immigration in the wage equation for the private service sector, with about same sized coefficients, as in the manufacturing sector wage equation.

The public sector wage equation shows contemporaneous effects of quarterly wage increases in both sector 1 and sector 2, but with more weight on the wage in the private sector. The wage relative to sector 2 is significant, which is consistent with the wage settlements in the public sector being last in the chain. We pick up a similar effect of the expected rise in the cost of living as in sector 2, which therefore emerges as a systemic feature of Norwegian wage formation. The same is true for the effect of the rate of unemployment. A permanent increase in the rate of unemployment lowers the relative wages in the sector, which is consistent with Figure 3.

The structural wage-setting model implies a restricted reduced form maximised likelihood that can be compared to the maximised likelihood of the unrestricted system, or VAR. Since we have a set of over-identifying restrictions, the statistical validity of the ordered wage-fixing system can be tested with respect to the unordered wage-setting system (the VAR) by a likelihood ratio test. This is the $\chi^2_{ENC-VAR}$ reported at the end of Table 9, and the interpretation of the p-value of 0.26 is that the ordered system represents no significant loss of explanatory power relative to the unrestricted and unordered VAR model. The unordered purely statistical system is encompassed by the structural model, see Hendry and Mizon (1993) and (Hendry 1995, Chapter 14). Bårdsen et al. (2005, Ch. 3) presents an application to Norwegian wage-price dynamics of this econometric approach.

We next turn to the detailed results for the dynamic specification of the 1980(1)-2000(4) dataset, cf. Table 10 before inflation targeting was formally introduced in 2001 and while labour immigration to Norway was at a low level. As noted in the main text, it is reasonable to include the unemployment rate and the immigration rate as unrestricted stationary variables on this sample, so that the equilibrium correction terms (denoted $\widehat{ecs}_{it}$ $i = 1, 2, 3$) become the wage share for sector 1, and the two relative wage rates for sectors 2 and 3.

Table 10: The dynamic specification of the leader-followership model in Table 8. Sample 1980(1) - 2000(4).

Sector 1 (manufacturing)
$\Delta w_{1t} = -\underset{(0.0345)}{0.1026}\,\Delta\tau_1 - \underset{(0.0498)}{0.27}\,\Delta_3 w_{1t-1} - \underset{(0.0196)}{0.04604}\,\widehat{ecs}_{1t-1} + \underset{(0.0622)}{0.3152}\,\Delta_4 p_t - \underset{(0.00488)}{0.01443}\,u_{t-1} - \underset{(0.0826)}{0.00003}\,IM_t$
Sector 2 (private services)
$\Delta w_{2t} = \underset{(0.09)}{0.1531}\,\Delta w_{1t} + \underset{(0.0675)}{0.2177}\,\Delta_3 w_{1t-1} - \underset{(0.0831)}{0.5042}\,\Delta_3 w_{2t-1} - \underset{(0.0949)}{0.1139}\,\widehat{ecs}_{2t-1} + \underset{(0.134)}{0.1079}\,\Delta_4 p_1 + \underset{(0.124)}{0.09343}\,\Delta_4 p_{t-1} - \underset{(0.00396)}{0.01391}\,u_{t-1} - \underset{(0.00569)}{0.01472}\,\Delta u_{t-2} - \underset{(0.0602)}{0.0314}\,IM_{t-4}$
Sector 3 (Public Sector)
$\Delta w_{3t} = \underset{(0.14)}{0.4935}\,\Delta w_{1t} + \underset{(0.14)}{0.4935}\,\Delta w_{2t} - \underset{(0.0582)}{0.03097}\,\Delta w_{3t-1} - \underset{(0.0581)}{0.3398}\,\Delta w_{3t-2} - \underset{(0.0484)}{0.1608}\,\Delta w_{3t-3} - \underset{(0.0823)}{0.2443}\,\widehat{ecs}_{3t-1} + \underset{(0.0425)}{0.1585}\,\Delta_4 p_t - \underset{(0.00296)}{0.01027}\,u_{t-1}$
cpi-equation:
$\Delta p_t = -\underset{(0.102)}{0.01002}\,\Delta p_{t-1} - \underset{(0.104)}{0.04616}\,\Delta p_{t-2} - \underset{(0.0884)}{0.2634}\,\Delta p_{t-3} + \underset{(0.0863)}{0.3841}\,\Delta p_{t-4} - \underset{(0.0184)}{0.07642}\,\widehat{ec}_{4t-1}$
misspecifications tests† and var encompassing test‡:
$F_{AR} = 1.1046[0.2890]$ $\chi^2_{NORM} = 14.251[0.0755]$ $\chi^2_{ENC-VAR} = 143.08[0.0000]$
Notes
Sample 1980(1)-2000(4). Estimation is by FIML. Deterministic terms are omitted Standard errors are in parentheses below the parameter estimates. $\widehat{ecs}_{it}$ $(i = 1, 2, 3)$ and $\widehat{ec}_{4t}$ are explained in the main text.: †System versions of 1-5 order autocorrelation and normality tests, see Doornik and Hendry (2013) ‡The likelihood-ratio test of over-identifying restrictions, see Doornik and Hendry (2013) The numbers in [ ] are p-values of these tests

The similarity of the results for the short-run wage leadership-followership sample compared to the full sample results is discussed in the main text. The implied long-run wage equations from the short-sample results become:

$$w_1 = \tau_t - x_{1t} - 0.31u - 0IM \tag{22}$$

$$w_2 = w_1 - 0.12u - 0.28IM \tag{23}$$

$$w_3 = w_2 - 0.04u \tag{24}$$

which are comparable to, e.g., Panel 4 in Table 4. The main difference is that the estimated effect of immigration on the wage-leading sector is zero on the short-sample, which

is reasonable since "normal" labour immigration to Norway probably has little impact on the bargaining power of the trade unions in the wage-leading sector. In the private wage-following sector, there is an estimated effect, which is also reasonable since there is more direct market regulation and weaker unions in this sector.

References

: n.d.

Aukrust, O.: 1977, Inflation in the open economy. A Norwegian model, *in* L. B. Klein and W. S. Sâlant (eds), *World Wide Inflation. Theory and Recent Experience*, Brookings, Washington D.C.

Bårdsen, G., Eitrheim, Ø., Jansen, E. S. and Nymoen, R.: 2005, *The Econometrics of Macroeconomic Modelling*, Oxford University Press, Oxford.

Bårdsen, G., Jansen, E. S. and Nymoen, R.: 2003, Econometric inflation targeting, *Econometrics Journal* **6**, 429—460.

Blanchard, O. and Galí, J.: 2010, Labor markets and monetary policy: A new keynesian model with unemployment., *American Economic Journal: Macroeconomics* **2**(2), 130.

Blanchard, O. J.: 1987, The wage price spiral, *Quarterly Journal of Economics* **101**, 543–565.

Blanchflower, D. G. and Oswald, A. J.: 1994, *The Wage Curve*, The MIT Press, Cambridge, Massachusetts.

Bowitz, E. and Cappelen, Å.: 2001, Modelling incomes polcies: Some norwegian experiences 1973-1993, *Economic Modelling* **18**, 349–379.

Bratsberg, B. and Raaum, O.: 2012, Immigration and wages: Evidence from construction, *The Economic Journal* **122**, 1177–1205.

Calmfors, L. and Seim, A. L.: 2013, Pattern bargaining and wage leadership in a small open economy, *Scandinavian Journal of Economics* **115**(1), 109–140. DOI: 10.1111/j.1467-9442.2012.01731.x.

Calvo, G.: 1983, Staggered prices in a utility maximizing framework, *Journal of Monetary Economics* **12**, 383–98.

Cuikerman, A. and Lippi, F.: 1999, Central bank independence, centralization of wage bargaining inflation and unemployment: Theory and some evidence, *European Economic Review* (43).

Doornik, J.: 2003, Asymptotic tables for cointegration tests based on the gamma-distribution approach. Nuffield College, University of Oxford.

Doornik, J. A.: 1998, Approximations to the asymptotic distributions of cointegration tests, *Journal of economic surverys* **12**(5), 573—593.

Doornik, J. A. and Hendry, D. F.: 2013, *Modelling Dynamic Systems PcGive 14. Volume 2*, Timberlake Consultants, London.

ECB: 2012, Euro area labour markets and the crisis, *Occasional Paper Series 138*, European Central Bank.

Edgren, G., Faxén, K. O. and Odher, C. E.: 1969, Wages, growth and distribution of income, *Swedish Journal of Economics* **71**, 133–60.

Forder, J.: 2014, *Macroeconomics and the Phillips Curve Myth*, Oxford Studies in the History of Economics, Oxford University Pres, Oxford.

Forslund, A., Gottfries, N. and Westermark, A.: 2008, Prices, productivity and wage bargaining in open economies, *Scandinavian Journal of Economics* **110**((1)), 169—195.

Gertler, M. and Trigari, A.: 2009, Unemployment fluctuations with staggered nash wage bargaining, *Journal of Political Economy* **177**(1), 3886.

Gjelsvik, M., Nymoen, R. and Sparrman, V.: 2013, Lnnsrelasjoner i kvarts og modag, *Technical Report 25*, Statistics Norway.

Hall, R. E.: 2005, Employment fluctuations with equilibrium wage stickiness., *American Economic Review* **95**(1), 50–65.

Harbo, I., Johansen, S., Nielsen, B. and Rahbek, A.: 1998, Asymptotic inference on cointegrating rank in partial system, *Journal of Business & Economic Statistics* **16**, 388–399.

Hendry, D. F.: 1995, *Dynamic Econometrics*, Oxford University Press, Oxford.

Hendry, D. F. and Doornik, J. A.: 2014, *Empirical Model Discovery and Theory Evaluation. Automatic Selection Methods in Econometrics*, Arne Ryde Memorial Lectures, MIT Press, Cambridge,Massachusetts.London, England.

Hendry, D. F. and Mizon, G. E.: 1993, Evaluating dynamic econometric models by encompassing the VAR, *in* P. C. B. Phillips (ed.), *Models, Methods and Applications of Econometrics*, Basil Blackwell, Oxford, pp. 272–300.

Hicks, J.: 1955, Economic foundations of wages policy, *Econimic Journal* **65**(259), 388–404.

Hoel, M. and Nymoen, R.: 1988, Wage formation in Norwegian manufacturing. An empirical application of a theoretical bargaining model, *European Economic Review* **32**, 977–997.

Holden, S.: 2005, Monetary regimes and the co-ordination of wage setting., *European Economic Review* (49), 833–843.

Johansen, K.: 1995a, Norwegian wage curves, *Oxford Bulletin of Economics and Statistics* **57**, 229—247.

Johansen, S.: 1995b, *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*, Oxford University Press, Oxford.

Johansen, S.: 2002, A small sample correction for the test of cointegrating rank in the vector autoregressive model, *Econometrica* **70**(5), 19291961. doi:10.1111/1468-0262.00358.

Krogh, T.: n.d., Real wage rigidity and the unemployment volatility puzzle in small open economies. Statistics Norway and Norges Bank.

Kwiatowski, D., Phillips, P., Schmidt, P. and Shin, Y.: 1992, Testing the null hypothesis of stationary against the alternative of a unit root, *Journal of Econometrics* (54), 159–178.

Layard, R., Nickell, S. and Jackman, R.: 1991, *Unemployment*, Oxford University Press, Oxford.

Marginson, P.: 2014, Coordinated bargaining in europe: From incremental corrosion to frontal assault, *European Journal of Industrial Relations* . Online version published 15 April 2014.

Meade, J.: 1982, *Wage-Fixing, Stagflation Volume 1*, George Allen and Unwin.

Nickell, S. J. and Andrews, M.: 1983, Unions, real-wages and employment in Britain 1951-79, *Oxford Economic Papers (Supplement)* **35**, 183–206.

Norges, B.: 2002, Inflation report, *Technical Report 3*, Norges Bank.

Nymoen, R.: 1989a, Modelling wages in the small open economy: An error-correction model of Norwegian manufacturing wages, *Oxford Bulletin of Economics and Statistics* **51**, 239–258.

Nymoen, R.: 1989b, Wages and the length of the working day. a empirical test based on Norwegian quarterly manufacturing data, *Scandinavian Journal of Economics* **91**, 599–612.

Nymoen, R.: 1991, A small linear model of wage- and price-inflation in the Norwegian economy, *Journal of Applied Econometrics* **6**, 255–269.

Nymoen, R. and Rødseth, A.: 1998, Modelling manufacturing wages in the nordic countries. Unpublished manuscript, University of Oslo.

OECD: 2012, Oecd economic outlook norway, *Technical report*, OECD.

OECD: 2014, *Sharing the pain equally? Wageadjustments during the crisis and recovery*, OECD Publishing, chapter 2. url = "/content/chapter/empl$_o utlook - 2014 - 5 - en$",.

Samuelson, P. A.: 1951, Economic theory and wages, *in* D. Wright (ed.), *The Impact of the Union*, Harcourt Brace & Company, pp. 312–43.

Solow, R. M.: 1990, *The Labour Market as a Social Institution*, Basil Blackwell, Oxford.

Soskice, D. and Iversen, T.: 2000, The nonneutrality of monetary policy with large price or wage setters, *Quarterly Journal of Economics* (1), 265–284.

Usher, D.: 2012, Bargaining unexplained, *Public Choice* **151**(1/2), 23–41.

Visser, J.: 2013, Wage bargaining institutions-from crisis to crisis, *European Economy. Economic Papers 488*, European Commision.

Have inflation targeting and EU labour immigration changed the system of wage formation in Norway?*

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Abstract

Collective agreements have played a central role in the system of wage formation in Norway for more than fifty years. Although the degree of coordination achieved has been variable, pattern wage bargaining has been a mainstay of the system. We investigate the degree of invariance in wage formation in Norway with respect to two recent structural changes: the transition towards inflation targeting in monetary policy and an unprecedented surge in labour supply due to higher immigration rates. We report empirical results that support the view that a semi-permanent high immigration may affect wages negatively in a significant way. However we do not find evidence that the stability of the arbitration system, and in particular the wage-bargaining pattern, has been changed by labour immigration or by inflation targeting monetary policy. An explanation of why we do not find evidence of structural changing effects of the transition of monetary policy, can be found in the fact that the wage arbitration system itself has syncronized the inflation expectations of the social partners. In that analysis, inflation targeting became a new layer of nominal stabilization, on top of the existing one.

Keywords: *Inflation modelling, pattern wage bargaining, inflation targeting, dynamic econometrics, cointegration, small open economy*

JEL classification: *C52, E24, E31, E37, J31.*

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1 Introduction

In Norway, the post-war system for wage formation was originally designed to balance the need for relative inflation control with the attainment of full employment and a balanced current account, cf. Aukrust (1977). The system has survived both the international stagflation of the 1970s and the Norwegian housing and banking crisis in 1989-90. It was based on an agreement between the government and the wage setters. The government committed to a fixed exchange rate policy, while the wage setters committed to collective wage bargaining. The system of wage fixing in Norway unfolds along both the horizontal and vertical dimension, implying that wage setting is based both on bargaining between evenly matched labour unions and firms' organizations, and bargaining at a high level. Later, similar systems were adopted in, for example, Sweden, see Edgren et al. (1969) and the Netherlands. A presentation of the Norwegian model from an econometric perspective is found in Bårdsen et al. (2005, Ch. 3).

The system for wage formation created a link between the wage norm for the overall economy, which was determined by the sector exposed to international competition, and profitability and productivity growth in the same sector. For several decades the system has linked wage increases in the other sectors of the economy (known as wage-followers) to wage formation in the exposed sector (the wage-leader). Down the years, it has been accepted, even supported, by governments of different party colours. Probably a recognition of the idea that this version of pattern bargaining represents an operational way of creating a certain degree of coordination in wage setting.

One threat to this system is the secular decline in the traditional manufacturing industries, which in Norway has been accelerated by the extraction of oil and gas and the use of the resulting 'oil-money'. Nevertheless, several commentators have noted the stability of the Norwegian norm-based system. For example by Visser (2013, p. 59-60), and also by OECD economists:

> ... the so called traditional ('non-oil') sector has been diminishing in importance. Despite this, it is important in wage setting. Rather than wages being determined by the relative bargaining strength of different sectors, the general wage level is set by the social partners first considering the wage increases that the traditional sector can "afford". OECD (2012, p. 15)

Nevertheless, the functioning of the wage formation system has been challenged by two important events in the last decade: the transition towards inflation targeting formally introduced in 2001 and the EU enlargement in 2004, which led to a large increase in the labour supply.

The change in monetary policy, from exchange rate stabilization to inflation targeting, can potentially change wage formation and the degree of coordination among wage setters, see Cuikerman and Lippi (1999), Soskice and Iversen (2000) and Holden (2005). In Norway, the introduction of inflation targeting in 2001 could potentially have a direct impact on wage formation, to the extent that it came into conflict with the 'existing order' based on manufacturing sector wage leadership. This potential for conflict was brought into focus by the central bank's *Inflation Report* from 2002. That report stated that the new monetary

policy required a reversal of the causality flow in wage formation.[1] The bank's worry was that the wage leader model could pull inflation in one direction (up for example, because productivity growth is higher there than in the service sectors), at the same time as monetary policy analysis suggested a reduction of aggregate demand relative to the bank's estimate of the Norwegian economy's potential output. This issue can also be related to theoretical research, showing that the pattern of wage bargaining under inflation targeting may be of marginal importance, e.g. Calmfors and Seim (2013). Hence, the hypothesis emerged that new monetary policy would force a change of roles in the pattern wage bargaining, where the wage-followers may become wage-leaders and vice versa.

With regards to immigration, research based on microeconomic data has found evidence that immigration has reduced wage earnings in certain groups in the labour market, see Bratsberg and Raaum (2012). The hypothesis is that, unless the immigrants emulate incumbent workers as regards trade union membership, increased immigration may reduce trade unions' bargaining power. The result might be that wage growth in sectors with high immigration rates deviates from the general wage growth.

We investigate empirically the degree of invariance in wage formation with respect to these structural changes. First, we investigate whether the value of labour productivity remains the main determinant of the wage trend in the exposed sector. Second, we test whether the wage norm created by the exposed sector still has a defining role in Norwegian wage formation. Third, we investigate whether inflation expectations have become more important after the introduction of inflation targeting. Finally, we include a labour immigration variable in our information set and identify its role in our structural model of wage formation.

We use quarterly data from 1980 to 2011. This means that about 1/3 of the sample comes from the era with a monetary policy regime that targeted inflation, and that about 1/4 of the sample comes from the recent period of increasing and high labour immigration. Hence, if either the monetary policy regime shift or the new immigration flow has affected wage formation or the social order with any force, the evidence should be in this data set.

The paper is organised as follows. In Section 2, we outline the theoretical framework of coordination in Norway and define the twin concepts of wage-leader and wage-follower sectors in wage formation. Based on the theoretical and statistical properties of the time series in Section 3, we formulate an empirical model, test the wage-leader/-follower properties, and carry out analysis of the model in Section 4. In Section 5, we conclude that the econometric results suggest that the Norwegian wage model has preserved its main properties in spite of the recent challenges to the system.

[1] The following quotation from Inflation Report number 3/2002, page 28 summarizes the hypothesis: *With an inflation target for monetary policy and a floating exchange rate, it is the inflation target, not wage growth abroad, that determines the level of growth which is consistent with stable profitability in the business sector over time. Inflation in Norway will over time be determined by the inflation target that the Government has set for monetary policy and not inflation abroad. Exchange rate developments are determined by inflation differentials between Norway and other countries.*

2 Wage formation

The twin concepts of wage-leader and wage-followers are important in the Norwegian system of wage formation. Wage-leader refers to the sector (or sectors) where the bargaining outcome defines the wage norm that forms the basis for bargaining in the other sectors of the economy (wage-followers). Throughout the whole post-war period, the settlements in the manufacturing sector have in practice defined the wage norm.[2]

We also define the manufacturing industry as the wage-leading sector but we test whether that role has been preserved so far in the new millennium. For the wage-followers, we have chosen to distinguish between the private service sector (including construction, mainly because of the importance of engineering consultancy in this sector) and the public sector. But also because negotiations in the public sector start after the wage settlements in private sector. Hence, in the following, we number the sectors 1 (wage-leader), 2 (wage-follower, private) and 3 (wage-follower, public).

2.1 The Wage-leader

The formulation of a theory of a sustainable wage norm requires an assessment of not only self-interest among workers and firms, but also of compromise. As pointed out by Usher (2012), 'compromise is then not just another way of talking about self-interest, and social, political and institutional forces are not merely cover-ups for imprecisely modelled individuals rational actions, they are among the fundamental determinants of decisions. In this view, even a full analysis of rational behaviour leads to an indeterminacy of wages, and other considerations had to be introduced to resolve it.

The recognition among economists that there is an indeterminacy in the economic theory of wages goes back to the 1950s, see Forder (2014, Ch. 1.4) who cites Samuelson (1951, p. 312) and Hicks (1955, p. 390) and other leading theorists. The economic theory of supply and demand could set some limits to what wages can be set, but within those limits closer determination requires that other relationships are introduced. The indeterminacy of wages from theory also characterizes the now standard Diamond-Mortensen-Pissarides (DMP) search and matching model. In the DMP model, the wage is usually determined in a Nash bargaining game. But is the wage logically equal to the Nash solution given the assumptions of the DMP model? As Hall (2005) pointed out, any wage in the bargaining set is in principle consistent with private efficiency on the part of both the firm and the worker. In that sense, the equilibrium wage rate is only set-identified. He then went on to analyze a solution where the real wage is fixed, which however is only one possibility of what in the DMP-literature is referred to as wage 'stickiness'. [3]

[2]Because of its importance, the system of wage-formation is a recurrent theme in public (and at times academic) debate. For example, the "traditional" definition of wage-leader has been challenged several times over the last decades. Those wanting to "change the content" of the wage-leading sector has emphasized both reduced legitimacy of the norm, as the importance of the sector for the total wage bill has become smaller, and reduced relevance (since Norwegian manufacturing may have become integrated in the super profitable oil and gas extraction). Among the arguments that are used in favour of no-change is that a manufacturing wage-norm aids coordination.

[3]Following Hall (2005), several papers have incorporated rigid wage setting in search models. For instance, Gertler and Trigari (2009) present a DMP model where the frequency of wage bargaining is constrained by

At the same time as we find it challenging to determine wages theoretically, we also observe that actual wage bargains are struck year after year, and that they are rationalized by considerations of profits, actual and required (to attract investments), cost of living and relative wages (fairness). These observed regularities give reason to believe that wage formation can be subject to econometric treatment. This is also what have motivated much of the econometric literature recently surveyed by Forder (2014). A development that goes back to the first half of the 1960s in Norway (although it was documented in English only later, by Aukrust (1977)) was that the wage norm became defined as a long-run trend, and also that short-run deviations had to be tolerated as 'part of the system'. But how should the trend be defined? Should it be a combination of the many things that might legitimate a secular trend trend? Or should the number of norm determined factors be limited? In Norway, a view that won support on both sides for the arbitration system, was that it was important to 'keep it simple' and relate the wage norm to the value of labour productivity in the exposed sector of the economy.

This idea of a wage norm can be treated econometrically using cointegration methods and equilibrium correction modelling. One of the main implications that we follow up is that, measured on a logarithmic scale, the wage level in the exposed sector, should be cointegrated with the log of the level of product prices and average labour productivity. Long-run price homogeneity together with a unit long-run elasticity on productivity lead to the implication of a stationary wage-share which may serve as an operationalization of an equilibrium wage-share. In the wider interpretation, the equilibrium income distribution needs not be completely constant but can depend on intermittent shifts, or trends, in bargaining power (possibly proxied by the unemployment rate) and the support to the idea about necessary compromise and coordination.

We base our econometric model on the assumption that both product price (q_1) and average labour productivity (z_1) are random-walk processes with drift. It follows that the time series for w_{1t} (the logarithm of the hourly wage in the wage-leader sector) contains both a deterministic and a stochastic (random walk) trend.[4]

Formally, the value of average labour productivity can be split into two components: q_{1t} (the log of the price in Norwegian currency) and z_{1t} (the log of average productivity in fixed prices), and these exogenous processes can be expressed as two random-walks with drift:

$$q_{1t} = q_{10} + q_{1t-1} + v_{q_1 t} \tag{1}$$

$$z_{1t} = z_{10} + z_{1t-1} + v_{a_1 t} \tag{2}$$

(1) and (2) imply that q_1 and z_1 are integrated of order 1, denoted $I(1)$. Positive deterministic trends in prices and productivity require $q_{10} > 0$ and $z_{10} > 0$.

Note that we could write equation (1) in terms of a price that is denoted in foreign currency, q^f_{1t}, and the exchange rate xr_t (log of kroner per currency unit) and specify time

Calvo (1983)) style lottery, leading to sticky wages. Blanchard and Galí (2010) combine a reduced form of search model with real wage rigidity with a New Keynesian model to study how this impacts monetary policy. Krogh (n.d.) generalizes the Hall-approach to a small open economy model where there is a non-trivial distinction between the consumer real wage and the producer real wage.

[4] In this section, we abstract from payroll tax rates in all three sectors. A representative payroll tax is included in the empirical model.

series models for q_{1t}^{f} and xr_t. To be consistent with $q_{10} > 0$ in equation (1), one or both of these time series models must contain a positive drift parameter. We assume that the price level variable q_{1t}^{f} has a positive drift, and for simplicity, we also assume that it is a constant parameter. The specification of the process for the nominal exchange rate is more complicated. For the period with a fixed exchange rate regime, it can possibly be specified without drift, but with intermittent structural breaks in order to represent devaluations. After the switch to the present floating exchange rate regime, a random walk, possibly with drift, is probably reasonable for the nominal exchange rate.

Hence, also across exchange rate regimes, equation (1) with $q_{10} > 0$ is a reasonable first approximation of the evolution of the price level in the wage-leading sector. As noted, there may be structural breaks in the drift parameter, and the foreign exchange rate regime-shift may also mean v_{q_1t} is not white-noise. The main point, however, is that the unit root property is robust across regimes and, in particular, that there are no explosive roots that logically must be included in the model of wage-price formation during the period with a floating exchange rate.

As noted, we assume that q_{1t} and z_{1t} jointly determine the deterministic trend in w_{1t}. In the simplest case, w_{1t} is cointegrated with the sum $q_{1t} + z_{1t}$, meaning that the logarithm of the wage share is I(0) with a constant mean s_1. This relationship can be written as:

$$w_{1t} - q_{1t} - z_{1t} = s_1 + e_{1t} \tag{3}$$

$$e_{1t} \sim I(0), \tag{4}$$

The sum of the productivity trend and the foreign price trend plays an important role in the framework, since it traces out a central tendency or a long-run sustainable scope for wage growth. This wage norm seems to correspond well to the concept of a wage corridor (Aukrust (1977)) for wage determination in the industries that are most exposed to foreign competition.

Cointegration implies equilibrium correction. Therefore, with reference to equations (1) and (2), e_{1t-1} should have significant predictive power for wage growth, Δw_{1t} ($\equiv w_{1t} - w_{1t-1}$):

$$\begin{array}{llcl} a) & (w_1 - q_1 - z_1)_{t-1} & \rightarrow & \Delta w_{1t} \\ b) & (w_2 - w_1)_{t-1} & \nrightarrow & \Delta w_{1t} \\ c) & (w_3 - w_1)_{t-1} & \nrightarrow & \Delta w_{1t} \end{array} \tag{5}$$

The economic interpretation is that collective wage bargaining, through its focus on the distribution of value added between labour and capital, implies that an equilibrium wage share (the parameter s_1) is maintained over time, cf. Forslund et al. (2008). (5b) and (5c) capture the idea that if the manufacturing sector is wage leading, it is implied that Δw_{1t} cannot be influenced by the lagged relative wage to the two other sectors.

As mentioned above, the equilibrium wage share is likely to depend on several underlying factors, e.g. related to production technology, product market conditions and bargaining power. It is only when these factors are constant that s_1 is likely to be a stable parameter. More generally, we can therefore add a time subscript to s_1 in equation (3) and write s_{1t} as a function of the variables that we include as its determinants in the econometric model, namely the log of the unemployment rate, u_t, and the immigration rate, IM_t:

$$s_{1t} = s_{10} + \beta_{11} u_t + \beta_{12} IM_t.$$

Hence, the extended long-run wage equation for sector 1 that we use in the following becomes

$$w_{1t} - q_{1t} - z_{1t} = s_{10} + \beta_{11} u_t + \beta_{12} IM_t + e_{1t} \tag{6}$$

If we abstract from IM, (7) is a standard linearised wage bargaining model, with full weight on producer prices and no weight on consumer prices, see e.g. Nickell and Andrews (1983) and Hoel and Nymoen (1988). The signs of the parameters, β_{11} and β_{12}, are expected to be negative (or zero). A higher unemployment rate can reduce the bargaining power of the unions. A marked change in the immigration rate may also affect union bargaining power and coordination negatively, unless a large proportion of the immigrants choose to become union members.

2.2 Wage-followers

If the manufacturing wage level represents the wage norm in the pattern of wage bargaining, and if $w_{1t} \sim I(1)$, w_{2t} must also be $I(1)$ and cointegrated with w_{1t}:

$$w_{2t} - w_{1t} = s_2 + e_{2t} \ , \tag{7}$$

$$e_{2t} \sim I(0), \tag{8}$$

where s_2 is the equilibrium relative wage. Note that equation (7) and (8) can be maintained by adjustments to wages in both sector 1 and sector 2. In order to define sector 2 as a wage-follower with sector 1 as a wage-leader we require:

$$\begin{array}{lccc} a) & \Delta w_{1t} & \overset{\longrightarrow}{\not\leftarrow} & \Delta w_{2t} \\ b) & (w_{2t-1} - w_{1t-1}) & \longrightarrow & \Delta w_{2t} \end{array} \tag{9}$$

Requirement (9a) implies that the contemporaneous relationship between the wage growth rates in the two sectors is recursive. We can implement and test (9a) as a restriction on a simultaneous equations model of wage setting. (9b) requires that the stationarity of the relative wage is due to equilibrium correction in sector 2.

In the Norwegian system of wage fixing, bargaining in the public sector starts after negotiations in the private sector, and it is therefore natural to assume that w_{3t} equilibrium corrects with respect to w_{2t}, i.e.:

$$w_{3t} - w_{2t} = s_3 + e_{3t} \tag{10}$$

$$e_{3t} \sim I(0). \tag{11}$$

The required recursive structure is the same as above, but between sector 2 and sector 3:

$$\begin{array}{lccc} a) & \Delta w_{2t} & \overset{\longrightarrow}{\not\leftarrow} & \Delta w_{3t} \\ b) & (w_{3t-1} - w_{2t-1}) & \longrightarrow & \Delta w_{3t} \end{array} \tag{12}$$

Clearly, for the government sector to be follower of sector 2, Δw_2 in (9) should not respond to $(w_3 - w_2)_{t-1}$. In both sectors, we include the possibility of potential wage shifting in the empirical model. Hence, s_2 in equation (7) and s_3 in equation (10) can include effects from the unemployment rate and from the immigration flow in the same way as in equation (6).

2.3 Wage-price inflation

The wage-price spiral is well known from the literature, cf. Blanchard (1987), Meade (1982) and Layard et al. (1991). In the wake of the financial and jobs crisis, the dynamics of wage and price setting have come to be regarded as central to rebalancing the euro area after the crisis, cf. ECB (2012) and OECD (2014).

Wage increases to compensate for expected increases in the cost of living are regularly demanded in the collective wage bargaining context, also in the wage-leading manufacturing sector. In Norway, a government-supported body, the Technical Calculation Committee for Wage Settlements (TCC), has the anchoring of CPI inflation forecasts as one of its tasks. The committee consists of representatives of both employer and employee organisations, and the consolidation of expectations takes place before the annual rounds of wage fixing starts.

Since CPI inflation, in turn, depends on growth in wage costs (in addition to the price of imported consumer goods), it is clear that a realistic system of equations for Δw_{1t}, Δw_{2t} and Δw_{3t} will not be completely recursive: in the short run, wage adjustments in the manufacturing sector also depend on CPI inflation expectations.

Nevertheless, as long as the wage norm that defines e_{1t} does not include the level of CPI, the inclusion of the rate of inflation in the system of wage formation represents a realistic modification of the otherwise recursive structure.

2.4 Potential reversal of bargaining pattern

As noted in the introduction, it is possible that inflation targeting caused an inversion of the causal ordering in wage formation. Specifically, that the relative wage $w_1 - w_2$ becomes a predictor of Δw_1, rather than of Δw_2. According to this hypothesis, neither (5b) nor (9b) will hold empirically on the latter sample period when monetary policy has changed. The hypotheses of reversed causality can be tested within the cointegrated VAR .The VAR is estimated in section 4.

3 Temporal properties of the time series variables

In this section, we first examine the assumption that wages, prices and productivity are non-stationary. Second, we examine whether the residuals of the equilibrium correction relationships are stationary. Stationary residuals between integrated variables imply a cointegrating relationship between the variables in the long-term solution.

Time series for wages, productivity and producer price developments are taken from the National Accounts. The unemployment rate is from the labour force survey, while consumer prices, immigration and population figures are collected from the official pages of Statistics Norway. Variables measured in natural logarithms are denoted by lower case Latin letters. Appendix A contains a detailed variable description.

3.1 The manufacturing sector

Wages, product prices and productivity in the manufacturing sector have increased over time, see Figure 1. The statistical properties of the variables are tested using two unit

root tests, the Augmented Dickey Fuller test (ADF) and a test by Kwiatowski et al. (1992) referred to as KPSS henceforth. The null of the ADF test is a unit root, and a rejection of the null implies that the time series is stationary. The null of the KPSS, on the other hand, is stationarity. Hence a rejection implies that the time series is integrated. Both tests are used to shed light on the time series properties.

The KPSS test rejects that w_1 is stationary, see Table 1. The ADF test statistic (including a constant, trend and seasonal dummies) is just above the critical value, which suggests the opposite result. The KPSS test result is supported by the testing of stationarity of the first difference of the same variable. The ADF test statistic of w_1 in difference (including a constant and seasonal dummies) is clearly above the critical value. Results from both ADF and KPSS tests indicate that the time series for product prices, q_1 is integrated. The ADF test statistic for productivity, z_1, does not reject I(1), while the KPSS test indicates that stationarity cannot be rejected. While the opposite is true for consumer prices, p_t, the ADF test rejects I(1), while the KPSS test strongly rejects stationarity. However, it is logically inconsistent to regard CPI inflation as I(1), while interpreting wages and import prices as stationary. In sum, we treat the variables as integrated of order 1 in the econometric modelling.

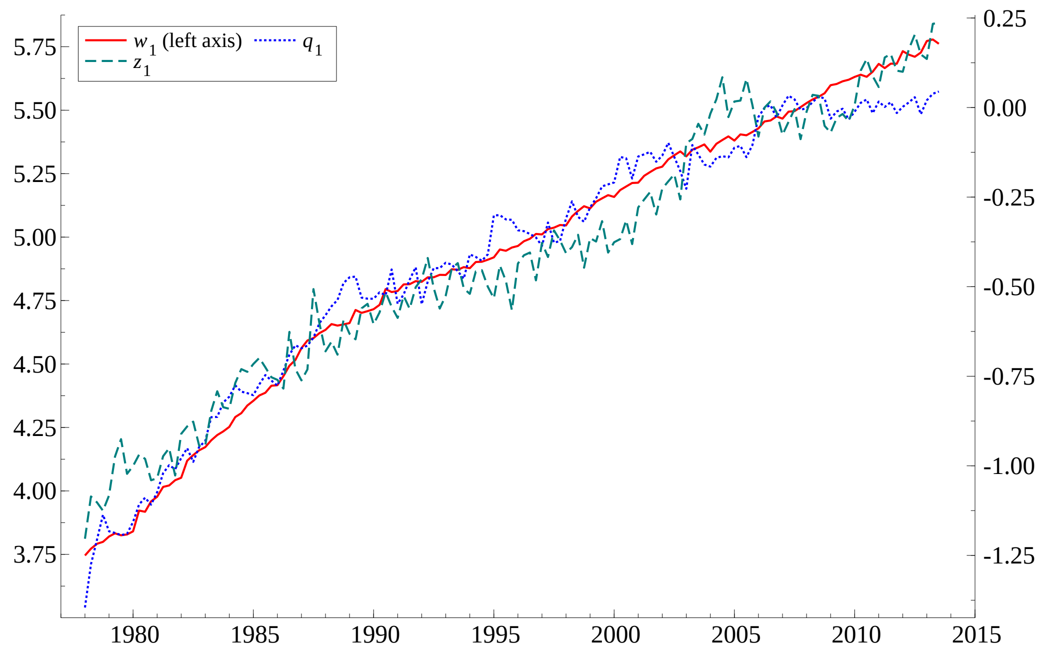

Figure 1: Components of the manufacturing wage share.

Figure 2 shows the unemployment rate, u_t, and the immigration (flow) rate , IM_t. Both variables are percentages (see Appendix A). The unemployment rate and the immigration rate both reject stationarity according to the KPSS test, see Table 1. The graph shows both the first period of departure from post-war full employment in the first half of the 1980s, and the return to near full employment in the years following deregulation of the financial markets (among other things). Next, unemployment increased again during the

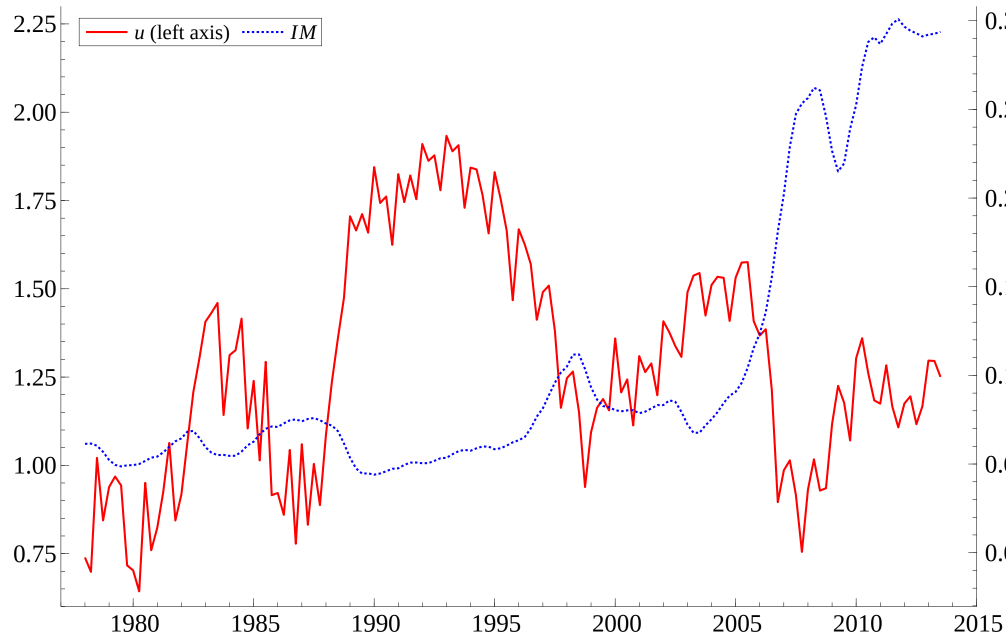

Figure 2: Plot of the logs of the unemployment rate and the immigration flow.

1990s following the fall in oil prices, a restrictive fiscal policy, the collapse of house prices and the ensuing banking crisis. Finally, the small changes in unemployment since 2005 are interesting. The number of employed wage earners in Norway grew by an unprecedented quarter of a million from 2005 to 2009. Naturally, the rate of unemployment fell, but not nearly as much as employment increased. This is due to the high immigration rates. This graph shows labour immigration to Norway from the EU and other developed countries. The migration inflow amounts to 1.6 per cent of the total labour supply in 2011 and is clearly large enough to significantly affect Norwegian labour markets and therefore possibly also wage setting.

We have also investigated the properties of the residuals from a regression of equation (3) and (6), see Table 2. The residuals from regression of equation (3) are labelled e_1, while the residuals from regression of equation (6) are labelled $e_{1,UIM}$. The KPSS test statistic for the wage share in Table 2 is lower and below the critical value if we control for shifts in labour immigration and unemployment. The tests therefore imply that the wage share is stationary when these variables are included. The ADF test does not reject the hypothesis of a unit root, neither for e_1 nor for $e_{1,UIM}$.

3.2 The public sector and private services

We have also investigated the statistical properties in the two other sectors; public and private services. Stationarity is rejected by the KPSS test for the hourly wage rate in both sectors. Figure 3 shows wages in sector 2 relative to sector 1, e_2, together with wages in sector 3 relative to sector 2, e_3. While wages in sector 2 appear to follow wages in sector 1 quite

Table 1: Unit root tests on the variables in levels and first differences. Sample period 1980(1) to 2011(4).[1]

	Level				First difference		
Variable	ADF (Lags)	KPSS	Char.[2]	Variable	ADF (Lags)	KPSS	Char.[2]
w_1	-3.714 (1)	0.282	c, t, s	Δw_1	-3.232 (4)	0.878	c, s
z_1	-2.255 (3)	0.076	c, t, s	Δz_1	-10.22 (2)	0.055	c, s
q_1	-2.382 (4)	0.263	c, t, s	Δq_1	-5.550 (3)	0.520	c, s
u	-2.810 (4)	0.205	c, t, s	Δu	-3.841 (3)	0.160	c, s
p	-5.979 (4)	0.310	c, t, s	Δp	-1.842 (3)	1.009	c, s
IM	-0.728 (2)	0.260	c, t, s	ΔIM	-5.352 (1)	0.401	c, s
w_2	-4.303 (4)	0.260	c, t, s	Δw_2	-2.659 (3)	0.818	c, s
w_3	-4.290 (0)	0.192	c, t, s	Δw_3	-9.680 (3)	0.760	c, s

[1]Note, however, that the sample period varies due to the number of lags in the ADF test.

[2]The characteristics are a constant (c), a trend (t) and seasonal dummies (s).

The 5% critical value of the ADF test is -3.45 and -2.88 in level and differences respectively.

The 5% critical value of KPSS test is 0.146 in level and 0.463 in differences.

closely, the wage level in sector 3 was markedly lower than the wage level in sector 2 for a long period. Although this reduction in the relative wage started in late 1980s, it is palusible that it was related to the rise in unemployment in the 1990s and the tight government budgets that characterised the first half of the decade, in particular. This indicates that, in the econometric model, the rate of unemployment can help to explain the development in relative wages. In the last period, wages in sector 3 have increased relative to wages in sector 2, while wages in sector 2 have decreased relative to sector 1. This is the period of high labour immigration.

In Table 2, the residuals from the regression of (7) are labelled e_2, while the residuals from a regression of (7), which includes u and IM, are labelled $e_{2,UI}$. The same regressions are performed for sector 3 as well, and residuals are labelled e_3 and $e_{3,UI}$, respectively. According to the ADF tests in Table 2, e_2, ec_{2UI} and ec_3 are non stationary. Again, however, the KPSS test supports the stationarity of the residuals and indicates that taking account of the possibility that unemployment and immigration can influence the long-run relationships, results in more stationary residuals.

Table 2: Unit root tests on the wage share and the relative wages. Sample period 1980(2) to 2011(4).

Variable	ADF (Lags)	KPSS	Char.[1]
e_1	-2.833 (3)	0.583	c
$e_{1,UIM}$	-2.751 (3)	0.305	c
e_2	-2.439 (3)	0.428	c
$e_{2,UIM}$	-2.232 (3)	0.155	c
e_3	-2.135 (3)	0.328	c
$e_{3,UIM}$	-3.153 (3)	0.170	c

The 5% critical value of the ADF test is -2.88 and 0.463 for the KPSS test
[1]The tests have included a constant (c) in the test.

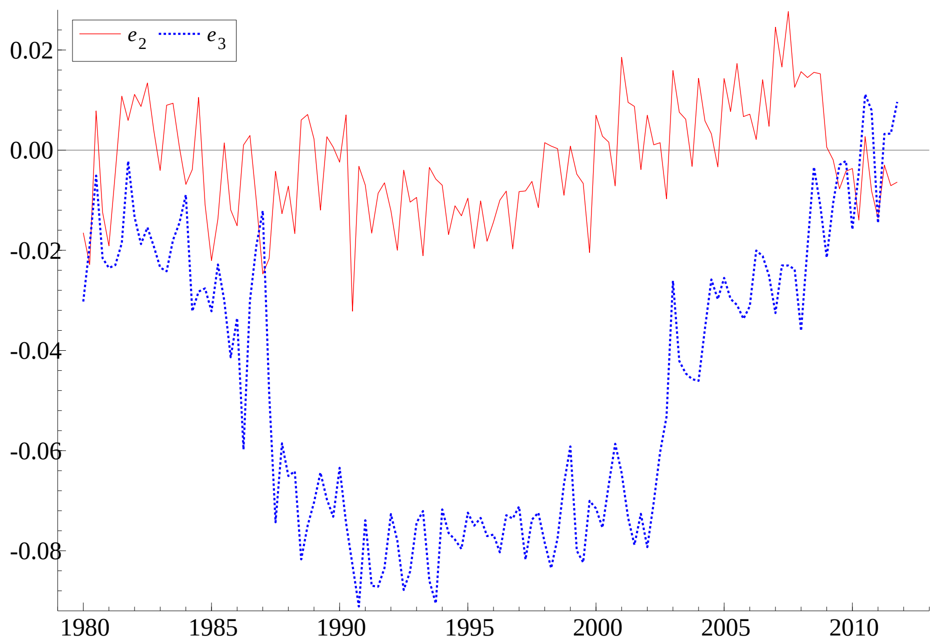

Figure 3: Wages in sector 2 relative to sector 1 (e_2) and wages in sector 3 relative to wages in sector 2 (e_3).

4 Econometric results

The previous section showed that univariate tests give clear support for unit-root non-stationarity, and some initial support for three cointegration relations of the pattern bargaining model. However, cointegration is a system property, and a concise way of testing the long-run theory is to formulate a vector autoregression (VAR) and apply the maximum likelihood method of Johansen (1995b). This step in the analysis is taken in section 4.1.

In section 4.2, we specify a cointegrated equilibrium correction model (ECM) that is consistent with the evidence from the VAR-based test of cointegration. The ECM is used to further test the hypotheses about the dynamics of wage leader- and follower bargaining model (the causal structure between the growth rates). The tests illustrate the relative invariance of the model parameters after the change from a fixed exchange rate to a floating exchange rate and the surge in immigration rates.

4.1 Cointegration

We formulate a fourth order VAR for the three wage levels, manufacturing (w_1), private services (w_2) and public sector (w_3). We include the endogenous variables of the VAR in the vector $\mathbf{Y_{wt}} = (w_{1t}, w_{2t}, w_{3t})'$. The VAR is expressed as:

$$\Delta \mathbf{Y_{wt}} = \sum_{i=1}^{3} \mathbf{\Gamma}_i \Delta \mathbf{Y_{wt-i}} + \mathbf{\Pi} \mathbf{Y_{t-1}} + \mathbf{\Upsilon} \mathbf{Q_t} + \varepsilon_{\mathbf{t}}, \tag{13}$$

where all variables in the cointegration space are defined in the extended data vector: $\mathbf{Y_t} = (\mathbf{Y_{wt}}, x_t, \tau_{1t}, u_t, IM_t)'$. $x = q_1 + z_1$ is the wage-scope variable in the manufacturing sector. The role of the payroll tax rate, τ_{1t}, enables us to distinguish between the wage rate per hour and the wage cost per hour. $\mathbf{\Gamma}_i$ $(i = 1, \ldots, 4)$ and $\mathbf{\Pi}$ are matrices with short-run and long-run parameters for wages, unemployment and immigration.

Formally, τ_{1t}, u_t and IM_t are also treated as I(1), since they are included in the $\mathbf{Y_t}$ vector. However, the tax rate is determined by political decisions and apparent unit root non-stationarity is probably due to policy determined changes in the mean. As the discussion above suggests, while u_t and IM_t are clearly non-stationary, the economically relevant 'model of stationarity' is probably that they are also driven by intermittent structural breaks in their means, rather than being convincingly represented by unit-root non-stationarity.

$\mathbf{Q_t}$ is other non-modelled variables in the VAR, namely the lagged annual CPI inflation rate and dummy variables. For simplicity, they are associated with the single matrix of parameters $\mathbf{\Upsilon}$. Finally, the vector $\varepsilon_{\mathbf{t}}$ contains the VAR residuals for the three wage growth rates.

As noted above, $\mathbf{Q_t}$ in (13) implies that the VAR includes other variables as well. In terms of economic interpretation, the most important variable is the lagged rate of consumer price inflation, which is included to control for the effects of cost of living considerations (as noted above). In the sample period, there have been periods of centralised wage settlements (wage laws), see e.g. Bowitz and Cappelen (2001) and Section 3. We include dummies for these events, and for other quarters with outliers. Finally, we include a constant, three centred seasonal dummies and a deterministic trend. The result is a relatively well-specified VAR. The residuals, $\hat{\varepsilon}_{\mathbf{t}}$, have a tendency to be negatively auto-correlated, indicating a certain over-fitting. However, the test of joint non-normality of $\varepsilon_{\mathbf{t}}$ is insignificant at the 5 % significance level.

$\mathbf{\Pi}$ is often referred to as the long-run matrix, and it has the dimensions 3×7. Let r denote the rank of $\mathbf{\Pi}$. Since x_t, τ_{1t}, u and IM are non-modelled variables, r can be 0, 1, 2 or 3. $r = 0$ means that there is no cointegration, which would imply rejection of the theory we

formulated above. If, on the other hand, $r = 3$, the variables x_t, τ_{1t}, u and IM represents four common trends in the three nominal wage rates. The theory of wage leadership points to the wage-scope variable $x_t = q_{1t} + z_{1t}$ as the most dominant trend in the system. This is because the evolution of the scope variable defines the upward trend in the wage corridor in the manufacturing sector. If wage leadership holds, this trend is transmitted to the two other wage rates.

$\mathbf{\Pi}$ can be written as $\mathbf{\Pi} = \boldsymbol{\alpha\beta}'$, where $\boldsymbol{\alpha}$ is $3 \times r$ and $\boldsymbol{\beta}$ ($7 \times r$) is the matrix with cointegration parameters. The theory presented in section 2 implies $\tau = 3$ with (3), (7) and (10) as the cointegration relationship.

When conducting the cointegration rank test a trend is included and restricted to lie in the cointegration space (i.e. included in $\mathbf{Y_{t-1}}$) such that critical value exist, see Harbo et al. (1998). Table 3, column *Trace test statistic* presents the test statistic of the cointegration analysis. The critical values of Doornik (2003) for zero, one and two cointegrating relationships are presented in Table 3, column five. The trace test statistic for two or fewer cointegration relationships is below the 10 % critical value for a standard integrated VAR reported in Doornik (2003). This implies that the test rejects two or fewer cointegration relationships at the 10 % significance level. This result supports the hypothesis of three cointegrating relationships. However, these critical values do not take into account exogenous variables. The critical values increase with the number of exogenous variables included, cf. Harbo et al. (1998). The critical values with four exogenous variables are shown in Table 3, last column.

The two first test statistics are well above their respective critical values (adjusted for exogenous variables), which supports that the number of stationary long-run relationships is at least two. The last row shows that we are unable to formally reject two, and accept three, cointegrating vectors. However, the critical values are for asymptotic distributions, and using these critical values may lead to over-rejecting in small samples, see Doornik (1998). The correction factor for critical values without exogenous variables is given in Johansen (2002). Correction factors do not exist for critical values in cointegration analyses with exogenous variables, but the correction factor without exogenous factors implies that the critical values of Harbo et al. (1998) should be interpreted with care. We conclude that cointegration tests, broadly interpreted, provide formal support for three cointegrating relationships ($r = 3$), but that equilibrium correction may be relatively weak.

Table 4, Panel 1, shows the maximised likelihood value, *Log L*, after we have omitted the deterministic trend from the three long-run relationships. The trend was included for the purpose of making correct inferences about the rank, but once the rank has been fixed to three, its significance can be tested by a standard Likelihood-Ratio (LR) test with a Chi-square distribution. In this case, the restriction is rejected at the 1 % level. However, keeping a trend is without meaning economically, since the wage share and relative wages will then drift deterministically. A closer inspection of the results also reveal that the trend is sensitive to if the first three years in public sector is included. There was a substantial wage stop in public sector which ended in the beginning of the 80s and wages increased quite rapidly in the following years. Hence, if the first three years are excluded the tests imply that the trend should be removed. We therefore conclude that the trend should be removed from the cointegration space. In Table 4, we use the result of a rank equal to 3, but use a VAR without trend when testing the identification of the cointegrating relationships.

Table 3: Tests of cointegration rank. Sample 1982(1)-2011(4).

Eigenvalue			*Trace*	10% *Critical value Doornik*	
(λ_i)	H_0	H_1	*Test statistic*	*No exogenous*	*With exogenous*
0.37	$r = 0$	$r >= 1$	116.5	39.7	60.5
0.23	$r <= 1$	$r >= 2$	51.3	23.3	37.7
0.17	$r <= 2$	$r >= 3$	14.3	10.7	18.5

Endogenous variables: wages in manufacturing (w_1), private services (w_2) and public sector (w_3).
Restricted variables: wage scope ($x = q_1 + z_1$) in manufacturing,
unemployment (u), immigration (IM) and a trend (t).
Unrestriced variables: payroll tax-rate, τ_1, constant, seasonal dummies and some specific dummies which capture important events in the Norwegian wage setting, see e.g. Bowitz and Cappelen (2001) and Section 4.

Removing the deterministic trend does not contribute to identification. In Panel 1, we therefore show the results after we have imposed identifying restrictions that are consistent with the wage leadership theory. Specifically, the payroll tax rate is only included in the cointegrating relationship, which is normalised on w_1. This separates the second and third relationships from the first. The exclusion of x from the second and third relationships provides more identifying information for these two relationships. In order to separate the second from the third relationship, we exclude w_1 and IM from the third, and w_3 from the second relationship. Finally, in order to identify the first relationship, both w_2 and w_3 are excluded from the relationship that is normalised on w_1. The identified long-run relationships are reported in Panel 1 of the table, together with the likelihood value and the Chi-square distributed test of the seven over-identifying restrictions that represent the wage leadership view of Norwegian wage formation. The restrictions are statistically acceptable at the 5 % level.

Table 4: Testing steady-state hypotheses

Unrestricted system for wages

$w_1 = \beta_{w2,1}w_2 + \beta_{w3,1}w_3 + \beta_{\tau,1}\tau + \beta_{x,1}x + \beta_{u,1}u + \beta_{IM,1}IM$

$w_2 = \beta_{w1,2}w_1 + \beta_{w3,2}w_3 + \beta_{\tau,3}\tau + \beta_{x,2}x + \beta_{u,2}u + \beta_{IM,2}IM$

$w_3 = \beta_{w1,3}w_1 + \beta_{w2,3}w_2 + \beta_{\tau,3}\tau + \beta_{x,3}x + \beta_{u,3}u + \beta_{IM,3}IM$

Unrestricted $Log\ L = 1365.79768$

Panel 1: Identified long-run relationships

$w_1 = -\underset{(.)}{1}\tau + \underset{(.)}{1}x \underset{(0.09)}{-0.38}u \underset{(0.42)}{-1.72}IM$

$w_2 = \underset{(.)}{1}w_1 \underset{(0.07)}{-0.40}u \underset{(0.20)}{-0.51}IM$

$w_3 = \underset{(.)}{1}w_2 \underset{(0.014)}{-0.07}u$

$Log\ L = 1359.27709$, $\chi^2(7) = 13.04[0.07]$

Panel 2: Manufacturing wage leadership, $\alpha_{w1,2} = \alpha_{w1,3} = 0$

$w_1 = -\underset{(.)}{1}\tau + \underset{(.)}{1}x \underset{(0.09)}{-0.41}u \underset{(0.43)}{-1.76}IM$

$w_2 = \underset{(.)}{1}w_1 \underset{(0.07)}{-0.40}u \underset{(0.21)}{-0.57}IM$

$w_3 = \underset{(.)}{1}w_2 \underset{(0.015)}{-0.08}u$

$Log\ L = 1359.25748$, $\chi^2(9) = 13.08[0.16]$

Additional restrictions: $\chi^2(2) = 0.04(0.98)$

Panel 3: No response of Δw_2 to $(w_{2t-1} - w_{3t-1})$, $\alpha_{w2,3} = 0$

$w_1 = -\underset{(.)}{1}\tau + \underset{(.)}{1}x \underset{(0.09)}{-0.41}u \underset{(0.43)}{-1.76}IM$

$w_2 = \underset{(.)}{1}w_1 \underset{(0.07)}{-0.42}u \underset{(0.19)}{-0.55}IM$

$w_3 = \underset{(.)}{1}w_2 \underset{(0.016)}{-0.08}u$

$Log\ L = 1359.18919$, $\chi^2(10) = 13.22[0.21]$

Additional restrictions: $\chi^2(1) = 0.14[0.71]$

Panel 4: No direct response of Δw_3 to manufacturing norm, $\alpha_{w3,1} = 0$

$w_1 = -\underset{(.)}{1}\tau + \underset{(.)}{1}x \underset{(0.12)}{-0.58}u \underset{(0.53)}{-2.45}IM$

$w_2 = \underset{(.)}{1}w_1 \underset{(0.03)}{-0.17}u \underset{(0.15)}{-0.18}IM$

$w_3 = \underset{(.)}{1}w_2 \underset{(0.01)}{-0.04}u$

$Log\ L = 1358.39232$, $\chi^2(11) = 14.81[0.19]$

Additional restrictions: $\chi^2(1) = 1.59[0.21]$

Conditional on the identified long-run relationships, we can begin to test for the minimum restrictions that imply no feed-back from w_2 and w_3 to w_1. The results are reported in Panel 2 of the table, using $\alpha_{i,j}$ to denote the element in row i, column j in the matrix with loadings parameters. The increase in the test of the over-identifying restrictions is small. Therefore it is not surprising that two zero restrictions that distinguish Panel 2 from Panel 1, are statistically acceptable (the incremental test yields $\chi^2(2) = 0.04[0.98]$). We note that the estimated cointegration parameters also change very little from Panel 1 to Panel 2.

We conclude that the tests in Panel 1 and 2 provide relatively clear support for the view that the wage-leading role of the manufacturing sector has been in operation during our sample period. The result is different from those estimated by Nymoen (1991), who was unable to impose the wage-leading restrictions when estimating on a sample that started in 1969(1) and ended in 1987(4). As a result, he specified a preferred model that allowed for wage-wage effects. The different samples might be relevant here, since the 1970s in particular were characterized by super high wage inflation in certain years. Some of this may have been driven by low coordination and wage-wage inflation.[5]

Returning to Table 4, Panel 3 and 4 test additional restrictions implied by pattern wage bargaining. In panel 3, we add the restriction that the deviations from the second and third cointegration relationships are not corrected by manufacturing wages. This result shows that the increase in the test of over-identifying restrictions is small, and the incremental test does not reject additional restrictions. The cointegration parameters are also very similar when we compare Panel 2 and Panel 3. Finally, in Panel 4, the restriction that wages in sector 3 does not equilibrium correct with respect to the first cointegrating relationship is imposed. Although this restriction is acceptable, we note that the parameter estimates change notably. We return to the question of a possible contemporaneous effect of w_1 on w_3 in the structural model below.

The point estimates in the best identified models, e.g. in Panel 3, deserve comment. In the manufacturing sector (Sector 1), the elasticity of wages ($\beta_{u,1}$) with respect to unemployment is higher in absolute value than the estimates in, e.g., Nymoen (1989) and Johansen (1995a), who used data from the 1980s.

This may be due to the larger variation of u_t in our sample, which includes the period after the Norwegian banking crisis in the 1990s. For Sector 2, the estimated unemployment elasticity is practically the same ($\hat{\beta}_{u,2} = -0.4$). We note, however, that this implies that the estimated long-run wage responsiveness for the private wage-following sector is twice as large as for the wage-leading sector. This is realistic, since this sector is characterized by weaker union organisation and lower capacity for multi-employer agreements than in the manufacturing sector. In the public sector, the estimated parameter of unemployment ($\beta_{u,3}$) is much lower (in absolute value) than in the two other sectors, which is also reasonable given that changes in overall unemployment have little direct relevance for the two bargaining parties.

Since labour immigration is a new phenomenon, there are no studies that can serve as direct references for discussing our results for the parameters $\beta_{IM,1}$ and $\beta_{IM,2}$ in Panel 3.

[5]There are other differences as well, both in the data definitions and in the econometric methodology. Nymoen did not identify the cointegrating relationships separately from the short-run dynamics, as we will do. This may have led to poorer identification of his long-run relationships.

However, one important study that use micro data reports evidence that immigration has causally reduced wage earnings in certain sectors of the Norwegian economy, see Bratsberg and Raaum (2012). Compared to our own earlier research using single equation estimation methods, the estimated value of $\hat{\beta}_{IM,1} = -1.7$ is very large, see Gjelsvik et al. (2013). Note, however, that the estimated standard error indicates a broad 95 % confidence interval. The single equation estimation just mentioned provides point estimates of somewhat less than -1 for the wage-leading sector, and this value is within the confidence interval with good margin.

Based on the above interpretation of the evidence we have specified the following equilibrium correction variables for use in a structural econometric model:

$$\widehat{ec}_{1t} = w_{1t} + \tau_t - x_t + 0.38u_t + 1.1IM_t \tag{14}$$

$$\widehat{ec}_{2t} = w_{2t} - w_{1t} + 0.40u_t + 0.51IM_t \tag{15}$$

$$\widehat{ec}_{3t} = w_{3t} - w_{2t} + 0.07u_t \tag{16}$$

The only significant difference from Panel 1, is that we have adjusted the point estimate of the coefficient of IM in (14). The estimation results of the dynamic structural model will confirm that this downsizing of the direct effect of immigration on the manufacturing wage is data-acceptable.

4.2 A dynamic simultaneous equations econometric model

In this section, we formulate and estimate a dynamic model model that is consistent with the cointegration analysis, and which is identified econometrically. At this stage in the analysis we include CPI inflation into the model in an explicit way, which was suggested to be important to wage formation in Section 2.

In the VAR above, the lagged four-quarter change in the log of the official consumer price index, denoted $\Delta_4 p_{t-1}$ was included among the variables in $\mathbf{Q_t}$ in (13). In the same way as for the lagged wage growth rates, we can think of lagged price inflation as a variable in a reduced form obtained from a simultaneous equations model, SEM, that includes price inflation among the endogenous variables. In order to model price growth jointly with wage growth, we need a separate long-run relationship for the price level. We do not carry out a cointegration analysis for the price level in this paper. Instead, we use the following equilibrium correction term

$$\widehat{ec}_{4t} = p_t - 0.6(w_{1t} + \tau_t - z_{1t}) - 0.4pi_i \tag{17}$$

where the elasticities for unit wage cost and import price, pi, are taken from the econometric results in Bårdsen et al. (2003).

The SEM model can formally be expressed as:

$$\mathbf{B_0 Y_{wpt}} = \sum_{i=1}^{3} \mathbf{B}_i \mathbf{Y_{wpt-i}} + \mathbf{AEC_{t-1}} + \Upsilon \mathbf{Q_t} + \epsilon_\mathbf{t}, \tag{18}$$

where $\mathbf{Y_{wpt}} = (\Delta w_{1t},\ \Delta w_{2t},\ \Delta w_{3t}, \Delta_4 p_t)'$ contains changes in wages and the consumer price index, $\mathbf{EC}$ contains the long run relationships in (14)-(17) and $\mathbf{Q_t}$ is redefined to exclude

consumer prices. $\mathbf{A}$ is a diagonal matrix and all the elements of the diagonal in $\mathbf{B_0}$ is equal to one.

The last four columns in table 5 show coefficient estimates ($\hat{\mathbf{A}}$) for an identified four equation SEM, where a diagonal matrix with adjustment coefficients for the four *ec*-terms is sufficient for identification on the rank condition.[6] The maximised log likelihood reported in the last row serves as the unrestricted likelihood value that we use to test the wage-leadership-followership restrictions in Table 6.[7] The two test statistics are (F_{AR}), for autocorrelation of order between 1 and 5, and non-normality (χ^2_{NORM}), calculated from the SEM residuals. They are system versions of autocorrelation and normality tests, see Doornik and Hendry (2013), and are reported with asymptotic p-values in square brackets.

Table 5: Wage-price SEM. FIML estimates of the contemporaneous coefficient matrix $\hat{\mathbf{B}}_\mathbf{0}$ (first four columns with estimates) and adjustment coefficients $\hat{\mathbf{A}}$ (last four columns). Standard errors in brackets below the estimates. Sample period 1980(1) to 2011(4).

	Δw_{1t}	Δw_{2t}	Δw_{3t}	$\Delta_4 p_t$	$\widehat{ec}_{1t-1}$	$\widehat{ec}_{2t-1}$	$\widehat{ec}_{3t-1}$	$\widehat{ec}_{4t-1}$
Δw_{1t}	-1	-0.05	-0.04	0.37	-0.04	0	0	0
		(0.15)	(0.10)	(0.06)	(0.01)			
Δw_{2t}	0.42	-1	0.03	0.05	0	-0.04	0	0
	(0.06)		(0.06)	(0.09)		(0.006)		
Δw_{3t}	0.17	0.24	-1	0.16	0	0	-0.19	0
	(0.08)	(0.12)		(0.03)			(0.03)	
$\Delta_4 p_t$	0.02	-0.12	0.09	-1	0	0	0	-0.10
	(0.06)	(0.12)	(0.09)					(0.02)

Log L= 1892.0816, $F_{AR} = 1.18[0.16]$ and $\chi^2_{NORM} = 12.6[0.12]$.

Table 5 shows that the diagonal elements of the matrix with adjustments coefficients for the *ec*-terms are statistically significant, although the numerical values are relatively small for the first two. This is consistent with the results in Table 3 and Table 4. The results for the contemporaneous parameters ($\hat{\mathbf{B}}_\mathbf{0}$) also show a clear pattern. In the row for Δw_{1t}, the two coefficients of Δw_{2t} and Δw_{3t} are negative ("wrong sign"), but they are statistically insignificant from zero values (judged by the standard errors). Conversely, Δw_{1t} has a sizeable estimated coefficient in the rows for Δw_{2t} and Δw_{3t}, which supports the hypothesis that the manufacturing sector is wage-leading, with private service production and the public sector as wage-followers.

The first row in Table 6 shows that three versions of the null hypothesis that manufacturing is *not* wage-leading, are formally rejected at the 5 % level. In the second line of tests, the second entry supports that the private service sector is wage-leading in relation to the public sector. The third row of tests shows that the coefficient of Δw_{3t} can be restricted to

[6]At this stage, there are no restrictions on the correlation matrix of the contemporaneous disturbances, which means that the rank condition is necessary and sufficient.

[7]Note that this likelihood is not comparable to the corresponding in Table 4, since the model in Table 5 includes CPI inflation as an endogenous variable

Table 6: Likelihood-ratio tests of wage-leader/follower restrictions on the model in Table 5

Restrictions:	Sec 1 $\nrightarrow$ Sec 2	Sec 1 $\nrightarrow$ Sec 3	Sec1 $\nrightarrow$ Sec 2 and 3
	$\chi^2(1) = 54.5^{**}$	$\chi^2(1) = 4.6^{*}$	$\chi^2(2) = 54.9^{**}$
Restrictions:	Sec 2 $\nrightarrow$ Sec 1	Sec 2 $\nrightarrow$ Sec 3	Sec 2 $\nrightarrow$ Sec 1 and 3
	$\chi^2(1) = 0.11$	$\chi^2(1) = 4.4^{*}$	$\chi^2(2) = 4.7$
Restrictions:	Sec 3 $\nrightarrow$ Sec 1	Sec 3 $\nrightarrow$ Sec 2	Sec 3 $\nrightarrow$ Sec 1 and 2
	$\chi^2(1) = 0.19$	$\chi^2(1) = 0.20$	$\chi^2(2) = 0.42$

* and **denotes significance at the 5% and 1 % levels.

zero in both the manufacturing sector wage equation and in the private service sector wage equation without any significant drop in the likelihood value.

We next turn to the role of price inflation as measured by $\Delta_4 p_t$. This variable is significant with a large estimated coefficient in the manufacturing wage equation. The coefficient of inflation in the equation for Δw_{2t} is close to zero, but can note that there is a significant effect of the lagged inflation rate in this equation, see Appendix B with detailed estimation results.

The last row of Table 5, with the results for the inflation equation $\Delta_4 p_t$, shows that there is little support for within-quarter effects of wage changes on inflation.[8] However, since $\widehat{ec}_{4t}$ includes wage costs (w_{1t}+ $\ln(1+\tau_t)$), we nevertheless have a closed feed-back loop between wages and CPI.

Table 7 shows the results of a restricted estimation where we have imposed all of the six exclusion restrictions discussed above. Compared to Table 5, the retained coefficients change very little.

Table 7: Wage-price SEM with wage leader-follower restrictions imposed. FIML estimates of $\hat{\mathbf{B}}_\mathbf{0}$ and $\hat{\mathbf{A}}$ in equation (18) with standard errors in brackets below the estimates. Sample period 1980(1) to 2011(4).

	Δw_{1t}	Δw_{2t}	Δw_{3t}	$\Delta_4 p_t$	$\widehat{ec}_{1t-1}$	$\widehat{ec}_{2t-1}$	$\widehat{ec}_{3t-1}$	$\widehat{ec}_{4t-1}$
Δw_{1t}	-1	0	0	0.35 (0.06)	-0.04 (0.01)	0	0	0
Δw_{2t}	0.42 (0.06)	-1	0	0.06 (0.09)	0	-0.04 (0.01)	0	0
Δw_{3t}	0.17 (0.08)	0.25 (0.12)	-1	0.16 (0.03)	0	0	-0.19 (0.03)	0
$\Delta_4 p_t$	0	0	0	-1	0	0	0	-0.10 (0.02)

Log L =1890.98642, $F_{AR} = 1.17[0.17]$ and $\chi^2_{NORM} = 13.0[0.11]$
LR test of wage-leader following restrictions: $\chi^2(6) = 2.19[0.14]$

[8]The estimated SEM includes three unrestricted lags in the quarterly inflation rate in the fourth row. This ensures that there are no restrictions implied by normalisation on $\Delta_4 p_t$ in the fourth row. This representation is convenient for modelling the effects of inflation on wage dynamics.

It is interesting that Table 5 and Table 7 show that inflation expectations[9] have a significant effect on wage formation in the manufacturing sector, and therefore on the wage norm of the system. Is this evidence of inflation targeting in the sense that the wage increases in the manufacturing sector are led by the inflation forecast of the Central Bank rather than by the evolution of the wage scope? In order to investigate this possibility, as well as the invariance of the system with respect to the monetary policy regime shift more generally, we turn to the estimation results on a sample that ends before the structural break in the market for foreign exchange and in interest rate determination.

4.3 Stability and invariance

In order to address the the question of stability of the wage formation structure, we have estimated the model on a sample that ends in 2000(4). This sample ends before the labour immigration from Europe started, and it is unlikely that the immigration effects on wages that we estimated above are invariant to the shortening of the sample.

Consequently, it does not make sense to define the three wage equilibrium correction terms in the same way as in (14)-(16). Instead, we remove IM_t and u_t from the equilibrium correction variables, which for the short sample are defined as:

$$\widehat{ecs}_{1t} = w_{1t} + \tau_t - x_{1t} \tag{19}$$

$$\widehat{ecs}_{2t} = w_{2t} - w_{1t} \tag{20}$$

$$\widehat{ecs}_{3t} = w_{3t} - w_{2t} \tag{21}$$

implying the same long-run relationships as in the full sample. The estimated system is therefore specified by including immigration and unemployment as exogenous I(0)-variables (but subject to breaks) in the system. For the CPI level, we use (17) as the equilibrium correction term also in the short sample.

The results of the estimation with the wage-leadership-follower restrictions imposed are shown in Table 8. The estimation results are very similar to the full sample estimates in Table 7, although the joint test of the six restrictions is marginally significant at the 5 % level. However, and as previously discussed, the test statistic is not corrected for the smaller sample size. The details show that this is mainly due to an estimated negative coefficient of Δw_{2t} in the Δw_{1t} equation, i.e. the same sign problem that we noted for the full sample results, but more pronounced on the short sample. Appendix B contains Table 10 with the detailed dynamic specification for the short sample.

[9] Inflation in period t is included as an endogenous variable in the model estimated by FIML. Therefore, the inflation variable in the manufacturing sector wage equation for example, can be interpreted as an expectations variable.

Table 8: Pre inflation targeting wage-price SEM with wage-leader-follower restrictions imposed. FIML estimates with standard errors in round brackets below the estimates. Sample period 1980(1) to 2000(4).

	Δw_{1t}	Δw_{2t}	Δw_{3t}	$\Delta_4 p_t$	$\widehat{ecs}_{1t-1}$	$\widehat{ecs}_{2t-1}$	$\widehat{ecs}_{3t-1}$	$\widehat{ec}_{4t-1}$
Δw_{1t}	-1	0	0	0.32 (0.06)	-0.05 (0.02)	0	0	0
Δw_{2t}	0.45 (0.06)	-1	0	0.11 (0.13)	0	-0.11 (0.09)	0	0
Δw_{3t}	0.15 (0.09)	0.49 (0.14)	-1	0.16 (0.04)	0	0	-0.24 (0.08)	0
$\Delta_4 p_t$	0	0	0	-1	0	0	0	-0.08 (0.02)

$Log\ L$ =1288.14658, $F_{AR} = 1.10[0.29]$ and $\chi^2_{NORM} = 14.3[0.08]$
LR test of wage-leader following restrictions: $\chi^2(6) = 12.39[0.05]$

The results show in particular that in manufacturing the estimated coefficient of $\Delta_4 p_t$ is 0.32, with a t-value of 5.1. This is very close to the full sample results. Taken together, the two estimates show that the effect of inflation expectations on the wage norm was in place before inflation targeting was introduced. This is not surprising since CPI expectations have been an important part of the wage setting system for decades. Coordination of inflation expectations among representatives of the employer and employee confederations has been one of the main purposes of the TCC institution long before inflation targeting was introduced.

Another way of illustrating the stability of the estimated structure is by forecasting. Figure 4 shows dynamic forecasts for the period 2001(1) to 2011(4). As can be expected, there are some examples of forecast failures (actuals outside the forecast uncertainty fans), in particular for inflation. However, a major structural break in the 40-quarter forecast period would have resulted in much clearer forecast.

Finally, we illustrate relative parameter stability in the dynamic multipliers of the structural model. In Figure 5, the impulse responses to a shock in the inflation equation (an inflation shock) are shown. There are two graphs in each panel, corresponding to the short sample estimation and the full sample estimation. In all panels, the dynamic multipliers are very similar, illustrating that the results of this policy analysis do not depend on the ten years of inflation targeting. The sign of the multipliers changes from positive to negative. The interpretation is that, because the long-run growth rate in the manufacturing sector depends on producer price growth and productivity growth, the short-run influence of a shock to inflation will mean overshooting in manufacturing sector wages. Because of the wage-leader role of manufacturing, the overshooting spills over to wages in the two other sectors.

Figure 6 shows the impulse response parameters of a shock to the immigration rate (IM). These impulse responses are not invariant. As noted above, this is not surprising given that the surge in immigration came after 2000 (cf. Figure 2). Specifically, it is the effects of increased immigration on wage formation in the manufacturing sector that are underestimated on the sample that ends in 2000(4). Since wage growth in the wage-leading sector is important for the overall nominal path of the Norwegian economy, this spills over

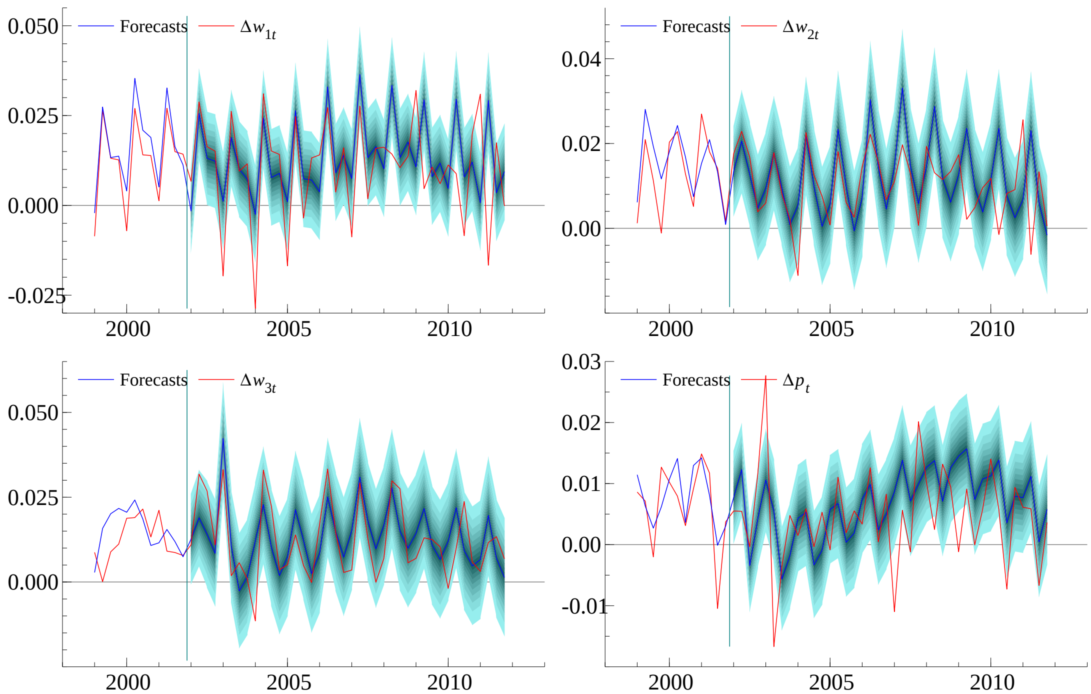

Figure 4: Dynamic forecasts with a 95 per cent forecast uncertainty fans. Forecasts are in blue, actuals in red. The estimation period is from 1980(1) to 2000(4).

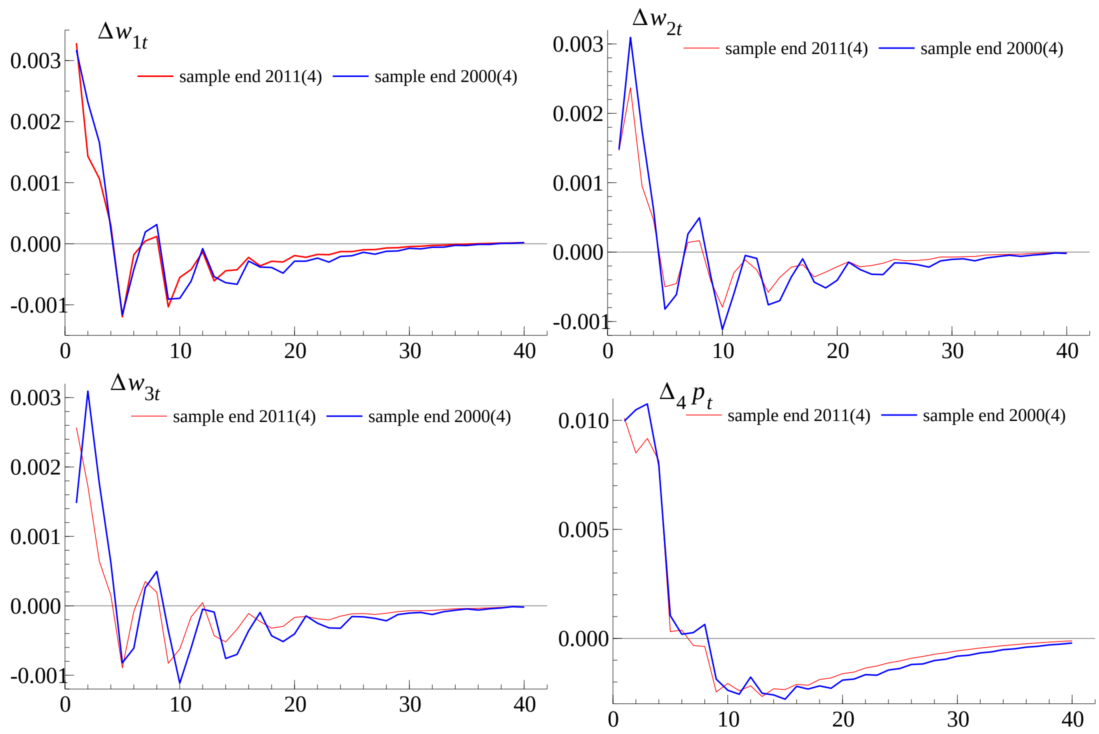

Figure 5: Impulse responses of a shock to the rate of quarterly inflation of 0.01 per cent.

to the inflation impulse responses in particular, which are severely underestimated on the sample that ends in 2000(4).

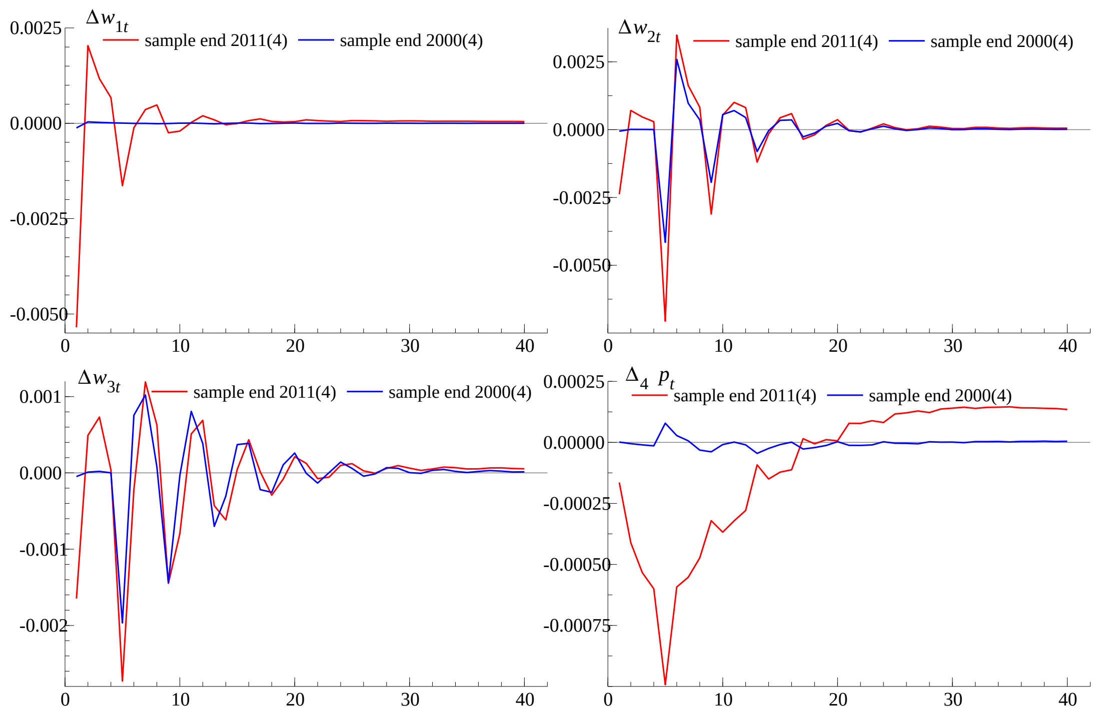

Figure 6: Impulse responses of a shock to the immigration rate by 0.10 percentage points.

5 Conclusion

In this paper we have modelled wages in three sectors of the Norwegian economy; manufacturing, private services and public sector. The model represents the typical pattern of wage bargaining in Norway, where manufacturing sector negotiates first and where the outcome is considered to be the wage norm for the other sectors of the economy. The logic of this pattern is simple, if the wage growth in the exposed sector becomes a wage norm for wage setting in the other sectors, the functional income distribution will be relative both in the wage-leading and wage-following sectors. The wage norm is defined in a system of arbitration that involves self-interest among both workers and firms, but, fundamentally also compromise. As noted in the introduction, the analysis of the system of wage formation as a 'social institution' applies quite, cf Solow (1990). generally. What is perhaps more special about Norway is the continued importance of collective wage agreements, and the roles that the confederate labour market organizations hold the tripartite interaction between government and social partners.[10].

Over the last 10-15 years, there have been two specific challenges to the Norwegian system of wage formation. First, the change to a floating exchange rate and, as a consequence, inflation targeting monetary policy. Second, the rise in labour immigration from outside Scandinavia. Both changes could potentially affect the bargaining power or the gain of coordinating wage formation, and hence change the recursive structure of wage formation in Norway.

The analytical framework that we specify, allows us to test the historical pattern of wage bargaining with a model with no pattern imposed. The model also allows the level of net immigration flow to affect wages separately.

The empirical results give the following answers to the questions posed in the introduction. The cointegration analysis show that there is a stable long term relationship between wage levels in the three sectors, with the manufacturing sector as the wage leader. Therefore, wage growth in manufacturing still constitute the wage norm of the Norwegian wage formation. The empirical results show that the wage norm in manufacturing depends heavily on the profitability growth in the same sector, which is consistent with the theory of collective wage bargaining, but the long term wage level is lower due to the negative and significant effect of the immigration flow. Immigration has also affected the relative wage between private and public services. This might suggest that immigration has affected bargaining power (a parameter of the system), but so far without fundamentally altering the pattern wage bargaining system. And although both unions and employment confederations have understood the nature and aims of monetary policy, we find no indication that the Norwegian monetary policy in 2001 has changed the system of wage formation or that inflation expectations have become more important.

Immigration effects apart, we find that the econometric model has more or less the same parameter estimates on two different samples: one with the period of inflation targeting, and another that ends before this policy change was formally introduced in 2001. The impulse responses of a shock to inflation also show a high degree of invariance. The recursive structure where the wage settlements in manufacturing represent a norm for wage setting is still an

[10]cf. OECD (2012, p. 15)

important part of the Norwegian wage settlements, despite both the structural break in monetary policy and high immigration rates. Hence, the corrosion of, or the attacks on the system of coordinated bargaining analyzed by Marginson (2014) have not been seen in Norway. Instead, the level of coordination is probably notably higher than it was during the 1980s, as also Visser (2013) concludes in his broad analysis of trends in wage bargaining institutions , see Visser (2013, p. 63)

One of the explanations why inflation targeting had little impact on the structure of wage setting, may be that the system of wage formation already included an important element of inflation forecasting and expectations formation. Formally, the Technical Calculation Committee for Wage Settlements has for decades helped to synchronize the inflation anticipations of unions and by firms. For this reason, the central bank's focus on inflation forecast and expectations when inflation targeting started in 2001, was not as "new" to wage and price setters as some commentators would have it.

There are of course other aspects of monetary policy that affect how difficult or easy it is to reconcile inflation targeting with the Norwegian system of wage formation. It is quite likely that a practice of strict inflation targeting (large weight on the inflation-gap and short policy horizon) would have put stress on the system. However, Norges Bank has presented itself as a super flexible inflation targeter. The indication is that the prospects for both unemployment and GDP growth carried large weights in the interest rate policy decisions already before the financial crisis in 2008. During most of the period of inflation targeting, Norges Bank have adopted a relatively long policy horizon, and has shown no haste in closing inflation gaps. The possibility that monetary policy in the future will become more strict to changes in wage setting remains only a vague prospect.

A Data definitions and sources

As explained in the text, lower case letters refer to the logarithm of the original variables listed below. For example, $u_t = \log(U_t)$ denotes the log of the unemployment rate. Variables in first differences are denoted by Δ. Subscripts denote time period. For example, p_{t-4} refers to the (log of) the price level four periods back.

$W_{it}-$Index for hourly wage in sector i=1,2,3
P_t- Consumer price index
$Q_{1t}-$ Price deflator of gross value added, manufacturing industry
$Z_{1t}-$ Labour productivity, output per hour in manufacturing
PI_t- Price deflator on imports of goods and services
U_t- Unemployment rate, in per cent. Civilian unemployment,
IM_t- Immigration from land group 1 and 2, in percent of the population aged 15-74. Group 1 include EU/EFTA countries, North America, Australia and New Zealand. Group 2 includes Eastern Europe except EU countries.
τ_t- represents the natural logarithm of the payroll tax rate plus one.

All variables are from the database of the macroeconometric model KVARTS, maintained by Statistics Norway.

B Additional estimation results

In this appendix, Table 9 shows the detailed dynamic specification of the model in Table 7, while Table 10 is the counterpart to Table 8 in the main text.

Table 9: The dynamic specification of the leader-followership model in Table 7

SECTOR 1 (MANUFACTURING)
$\Delta w_{1t} = -\underset{(0.0368)}{0.1198}\ \Delta \ln(1+\tau_t) - \underset{(0.0489)}{0.3834}\ \Delta_3 w_{1t-1} - \underset{(0.00797)}{0.04055}\ \widehat{ec}_{1t-1} + \underset{(0.042)}{0.3462}\ \Delta_4 p_t$
SECTOR 2 (PRIVATE SERVICE)
$\Delta w_{2t} = \underset{(0.0558)}{0.4245}\ \Delta w_{1t} + \underset{(0.0458)}{0.2073}\ \Delta_3 w_{1t-1} - \underset{(0.054)}{0.5011}\ \Delta_3 w_{2t-1} - \underset{(0.00553)}{0.03664}\ \widehat{ec}_{2t-1}$ $+ \underset{(0.0853)}{0.07038}\ \Delta_4 p_t + \underset{(0.0792)}{0.1422}\ \Delta_4 p_{t-1} - \underset{(0.00476)}{0.01784}\ \Delta u_{t-2}$
SECTOR 3 (PUBLIC SECTOR)
$\Delta w_{3t} = \underset{(0.0787)}{0.1659}\ \Delta w_{1t} + \underset{(0.117)}{0.2506}\ \Delta w_{2t} - \underset{(0.0467)}{0.08951}\ \Delta w_{3t-1} - \underset{(0.0502)}{0.3811}\ \Delta w_{3t-2}$ $- \underset{(0.0454)}{0.2221}\ \Delta w_{3t-3} - \underset{(0.0311)}{0.1857}\ \widehat{ec}_{3t-1} + \underset{(0.0322)}{0.1578}\ \Delta_4 p_t$
CPI-EQUATION:
$\Delta p_t = -\underset{(0.0807)}{0.1779}\ \Delta p_{t-1} - \underset{(0.0766)}{0.02421}\ \Delta p_{t-2} + \underset{(0.0727)}{0.129}\ \Delta p_{t-3} + \underset{(0.0736)}{0.2619}\ \Delta p_{t-4}$ $- \underset{(0.0166)}{0.1047}\ \widehat{ec}_{4t-1}$
MIS-SPECIFICATIONS TESTS[†] AND ENCOMPASSING (THE VAR) TEST[‡]:
$F_{AR} = 1.1666[0.1748]$ $\chi^2_{NORM} = 13.001[0.1118]$ $\chi^2_{ENC-VAR} = 83.025[0.0653]$
Notes
Sample 1980(1)-2011(4). Estimation is by FIML. Deterministic terms are omitted Standard errors are in parentheses below the parameter estimates. $\widehat{ec}_{i_{t-1}}$ $(i = 1, 2, 3, 4)$ are explained in the text. [†]System versions of 1-5 order autocorrelation test and normality test, see Doornik and Hendry (2013) [†]The likelihood-ratio test of the over-identifying restriction, see Doornik and Hendry (2013) The numbers in [ ] are p-values of these tests

The quarterly rate of change in the manufacturing wage rate is seen to depend negatively on the change in the payroll tax rate. This implies that tax increases are rolled back to the wage earners. Since $\widehat{ec}_{1t}$ includes $w_{1t} + \tau_t - q_{1t} - z_{1t}$, the wage level increases one-for-one with a reduction in the payroll tax rate. This means that the hourly wage cost is independent of the payroll tax rate in the long run. This result has been found in earlier empirical studies, also on aggregated data. The negative coefficient of the three-quarter lagged change in the wage rate ($\Delta_3 w_{1t-1}$) is also known from earlier studies, see for example Nymoen (1989a). It reflects the fact that most of the adjustment of the wage level takes place in the same quarter

each year (the second). The estimated effect of an expected rise in the annual inflation rate ($\Delta_4 p_t$) of 1 percentage point is as high as 0.3462. The effect is strongly significant with an implied t-value of 8.08. The implication for the wage norm is discussed in the main text.

The two last variables in the manufacturing wage equation are the unemployment rate (u_t), in log form, and net labour immigration (IM_t) in per cent. The unemployment effect is well known from previous studies and represents a moderately convex wage curve, see e.g. Blanchflower and Oswald (1994), Hoel and Nymoen (1988) and Nymoen and Rødseth (1998). The effect of IM_t could not have been estimated on samples that do not include the massive inflow of labour immigrants since 2005.

Wage growth in the private service sector is strongly influenced by both the (expected) wage norm (Δw_{1t}) and by lagged manufacturing wage increases ($\Delta_3 w_{1t-1}$). The coefficient of the lagged relative wage with respect to manufacturing ($\hat{ec}_{2t-1}$) is also sizeable and significantly different from zero. Taken together, this is strong evidence in favour of wage-following behaviour. The finding that the annual CPI inflation rate enters at one lag, and with a lower estimated parameter than in manufacturing, also indicates that wage adjustments in the private service sector are mainly anchored by the wage norm. That said, we also estimate significant effects of unemployment and immigration in the wage equation for the private service sector, with about same sized coefficients, as in the manufacturing sector wage equation.

The public sector wage equation shows contemporaneous effects of quarterly wage increases in both sector 1 and sector 2, but with more weight on the wage in the private sector. The wage relative to sector 2 is significant, which is consistent with the wage settlements in the public sector being last in the chain. We pick up a similar effect of the expected rise in the cost of living as in sector 2, which therefore emerges as a systemic feature of Norwegian wage formation. The same is true for the effect of the rate of unemployment. A permanent increase in the rate of unemployment lowers the relative wages in the sector, which is consistent with Figure 3.

The structural wage-setting model implies a restricted reduced form maximised likelihood that can be compared to the maximised likelihood of the unrestricted system, or VAR. Since we have a set of over-identifying restrictions, the statistical validity of the ordered wage-fixing system can be tested with respect to the unordered wage-setting system (the VAR) by a likelihood ratio test. This is the $\chi^2_{ENC-VAR}$ reported at the end of Table 9, and the interpretation of the p-value of 0.26 is that the ordered system represents no significant loss of explanatory power relative to the unrestricted and unordered VAR model. The unordered purely statistical system is encompassed by the structural model, see Hendry and Mizon (1993) and (Hendry 1995, Chapter 14). Bårdsen et al. (2005, Ch. 3) presents an application to Norwegian wage-price dynamics of this econometric approach.

We next turn to the detailed results for the dynamic specification of the 1980(1)-2000(4) dataset, cf. Table 10 before inflation targeting was formally introduced in 2001 and while labour immigration to Norway was at a low level. As noted in the main text, it is reasonable to include the unemployment rate and the immigration rate as unrestricted stationary variables on this sample, so that the equilibrium correction terms (denoted $\widehat{ecs}_{it}$ $i = 1, 2, 3$) become the wage share for sector 1, and the two relative wage rates for sectors 2 and 3.

Table 10: The dynamic specification of the leader-followership model in Table 8. Sample 1980(1) - 2000(4).

SECTOR 1 (MANUFACTURING)
$\Delta w_{1t} = \underset{(0.0345)}{-0.1026}\ \Delta\tau_1 - \underset{(0.0498)}{0.27}\ \Delta_3 w_{1t-1} - \underset{(0.0196)}{0.04604}\ \widehat{ecs}_{1t-1} + \underset{(0.0622)}{0.3152}\ \Delta_4 p_t - \underset{(0.00488)}{0.01443}\ u_{t-1} - \underset{(0.0826)}{0.00003}\ IM_t$
SECTOR 2 (PRIVATE SERVICES)
$\Delta w_{2t} = \underset{(0.09)}{0.1531}\ \Delta w_{1t} + \underset{(0.0675)}{0.2177}\ \Delta_3 w_{1t-1} - \underset{(0.0831)}{0.5042}\ \Delta_3 w_{2t-1} - \underset{(0.0949)}{0.1139}\ \widehat{ecs}_{2t-1} + \underset{(0.134)}{0.1079}\ \Delta_4 p_1 + \underset{(0.124)}{0.09343}\ \Delta_4 p_{t-1} - \underset{(0.00396)}{0.01391}\ u_{t-1} - \underset{(0.00569)}{0.01472}\ \Delta u_{t-2} - \underset{(0.0602)}{0.0314}\ IM_{t-4}$
SECTOR 3 (PUBLIC SECTOR)
$\Delta w_{3t} = \underset{(0.14)}{0.4935}\ \Delta w_{1t} + \underset{(0.14)}{0.4935}\ \Delta w_{2t} - \underset{(0.0582)}{0.03097}\ \Delta w_{3t-1} - \underset{(0.0581)}{0.3398}\ \Delta w_{3t-2} - \underset{(0.0484)}{0.1608}\ \Delta w_{3t-3} - \underset{(0.0823)}{0.2443}\ \widehat{ecs}_{3t-1} + \underset{(0.0425)}{0.1585}\ \Delta_4 p_t - \underset{(0.00296)}{0.01027}\ u_{t-1}$
CPI-EQUATION:
$\Delta p_t = \underset{(0.102)}{-0.01002}\ \Delta p_{t-1} - \underset{(0.104)}{0.04616}\ \Delta p_{t-2} - \underset{(0.0884)}{0.2634}\ \Delta p_{t-3} + \underset{(0.0863)}{0.3841}\ \Delta p_{t-4} - \underset{(0.0184)}{0.07642}\ \widehat{ec}_{4t-1}$
MISSPECIFICATIONS TESTS† AND VAR ENCOMPASSING TEST‡:
$F_{AR} = 1.1046[0.2890]$ $\chi^2_{NORM} = 14.251[0.0755]$ $\chi^2_{ENC-VAR} = 143.08[0.0000]$
Notes
Sample 1980(1)-2000(4). Estimation is by FIML. Deterministic terms are omitted Standard errors are in parentheses below the parameter estimates. $\widehat{ecs}_{it}$ $(i = 1, 2, 3)$ and $\widehat{ec}_{4t}$ are explained in the main text.: †System versions of 1-5 order autocorrelation and normality tests, see Doornik and Hendry (2013) ‡The likelihood-ratio test of over-identifying restrictions, see Doornik and Hendry (2013) The numbers in [ ] are p-values of these tests

The similarity of the results for the short-run wage leadership-followership sample compared to the full sample results is discussed in the main text. The implied long-run wage equations from the short-sample results become:

$$w_1 = \tau_t - x_{1t} - 0.31u - 0IM \quad (22)$$

$$w_2 = w_1 - 0.12u - 0.28IM \quad (23)$$

$$w_3 = w_2 - 0.04u \quad (24)$$

which are comparable to, e.g., Panel 4 in Table 4. The main difference is that the estimated effect of immigration on the wage-leading sector is zero on the short-sample, which

is reasonable since “normal” labour immigration to Norway probably has little impact on the bargaining power of the trade unions in the wage-leading sector. In the private wage-following sector, there is an estimated effect, which is also reasonable since there is more direct market regulation and weaker unions in this sector.

References

: n.d.

Aukrust, O.: 1977, Inflation in the open economy. A Norwegian model, *in* L. B. Klein and W. S. Sâlant (eds), *World Wide Inflation. Theory and Recent Experience*, Brookings, Washington D.C.

Bårdsen, G., Eitrheim, Ø., Jansen, E. S. and Nymoen, R.: 2005, *The Econometrics of Macroeconomic Modelling*, Oxford University Press, Oxford.

Bårdsen, G., Jansen, E. S. and Nymoen, R.: 2003, Econometric inflation targeting, *Econometrics Journal* **6**, 429—460.

Blanchard, O. and Galí, J.: 2010, Labor markets and monetary policy: A new keynesian model with unemployment., *American Economic Journal: Macroeconomics* **2**(2), 130.

Blanchard, O. J.: 1987, The wage price spiral, *Quarterly Journal of Economics* **101**, 543–565.

Blanchflower, D. G. and Oswald, A. J.: 1994, *The Wage Curve*, The MIT Press, Cambridge, Massachusetts.

Bowitz, E. and Cappelen, Å.: 2001, Modelling incomes polcies: Some norwegian experiences 1973-1993, *Economic Modelling* **18**, 349–379.

Bratsberg, B. and Raaum, O.: 2012, Immigration and wages: Evidence from construction, *The Economic Journal* **122**, 1177–1205.

Calmfors, L. and Seim, A. L.: 2013, Pattern bargaining and wage leadership in a small open economy, *Scandinavian Journal of Economics* **115**(1), 109–140. DOI: 10.1111/j.1467-9442.2012.01731.x.

Calvo, G.: 1983, Staggered prices in a utility maximizing framework, *Journal of Monetary Economics* **12**, 383–98.

Cuikerman, A. and Lippi, F.: 1999, Central bank independence, centralization of wage bargaining inflation and unemployment: Theory and some evidence, *European Economic Review* (43).

Doornik, J.: 2003, Asymptotic tables for cointegration tests based on the gamma-distribution approach. Nuffield College, University of Oxford.

Doornik, J. A.: 1998, Approximations to the asymptotic distributions of cointegration tests, *Journal of economic surverys* **12**(5), 573—593.

Doornik, J. A. and Hendry, D. F.: 2013, *Modelling Dynamic Systems PcGive 14. Volume 2*, Timberlake Consultants, London.

ECB: 2012, Euro area labour markets and the crisis, *Occasional Paper Series 138*, European Central Bank.

Edgren, G., Faxén, K. O. and Odher, C. E.: 1969, Wages, growth and distribution of income, *Swedish Journal of Economics* **71**, 133–60.

Forder, J.: 2014, *Macroeconomics and the Phillips Curve Myth*, Oxford Studies in the History of Economics, Oxford University Pres, Oxford.

Forslund, A., Gottfries, N. and Westermark, A.: 2008, Prices, productivity and wage bargaining in open economies, *Scandinavian Journal of Economics* **110**((1)), 169—195.

Gertler, M. and Trigari, A.: 2009, Unemployment fluctuations with staggered nash wage bargaining, *Journal of Political Economy* **177**(1), 3886.

Gjelsvik, M., Nymoen, R. and Sparrman, V.: 2013, Lnnsrelasjoner i kvarts og modag, *Technical Report 25*, Statistics Norway.

Hall, R. E.: 2005, Employment fluctuations with equilibrium wage stickiness., *American Economic Review* **95**(1), 50–65.

Harbo, I., Johansen, S., Nielsen, B. and Rahbek, A.: 1998, Asymptotic inference on cointegrating rank in partial system, *Journal of Business & Economic Statistics* **16**, 388–399.

Hendry, D. F.: 1995, *Dynamic Econometrics*, Oxford University Press, Oxford.

Hendry, D. F. and Doornik, J. A.: 2014, *Empirical Model Discovery and Theory Evaluation. Automatic Selection Methods in Econometrics*, Arne Ryde Memorial Lectures, MIT Press, Cambridge,Massachusetts.London, England.

Hendry, D. F. and Mizon, G. E.: 1993, Evaluating dynamic econometric models by encompassing the VAR, *in* P. C. B. Phillips (ed.), *Models, Methods and Applications of Econometrics*, Basil Blackwell, Oxford, pp. 272–300.

Hicks, J.: 1955, Economic foundations of wages policy, *Econimic Journal* **65**(259), 388–404.

Hoel, M. and Nymoen, R.: 1988, Wage formation in Norwegian manufacturing. An empirical application of a theoretical bargaining model, *European Economic Review* **32**, 977–997.

Holden, S.: 2005, Monetary regimes and the co-ordination of wage setting., *European Economic Review* (49), 833–843.

Johansen, K.: 1995a, Norwegian wage curves, *Oxford Bulletin of Economics and Statistics* **57**, 229—247.

Johansen, S.: 1995b, *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*, Oxford University Press, Oxford.

Johansen, S.: 2002, A small sample correction for the test of cointegrating rank in the vector autoregressive model, *Econometrica* **70**(5), 19291961. doi:10.1111/1468-0262.00358.

Krogh, T.: n.d., Real wage rigidity and the unemployment volatility puzzle in small open economies. Statistics Norway and Norges Bank.

Kwiatowski, D., Phillips, P., Schmidt, P. and Shin, Y.: 1992, Testing the null hypothesis of stationary against the alternative of a unit root, *Journal of Econometrics* (54), 159–178.

Layard, R., Nickell, S. and Jackman, R.: 1991, *Unemployment*, Oxford University Press, Oxford.

Marginson, P.: 2014, Coordinated bargaining in europe: From incremental corrosion to frontal assault, *European Journal of Industrial Relations* . Online version published 15 April 2014.

Meade, J.: 1982, *Wage-Fixing, Stagflation Volume 1*, George Allen and Unwin.

Nickell, S. J. and Andrews, M.: 1983, Unions, real-wages and employment in Britain 1951-79, *Oxford Economic Papers (Supplement)* **35**, 183–206.

Norges, B.: 2002, Inflation report, *Technical Report 3*, Norges Bank.

Nymoen, R.: 1989a, Modelling wages in the small open economy: An error-correction model of Norwegian manufacturing wages, *Oxford Bulletin of Economics and Statistics* **51**, 239–258.

Nymoen, R.: 1989b, Wages and the length of the working day. a empirical test based on Norwegian quarterly manufacturing data, *Scandinavian Journal of Economics* **91**, 599–612.

Nymoen, R.: 1991, A small linear model of wage- and price-inflation in the Norwegian economy, *Journal of Applied Econometrics* **6**, 255–269.

Nymoen, R. and Rødseth, A.: 1998, Modelling manufacturing wages in the nordic countries. Unpublished manuscript, University of Oslo.

OECD: 2012, Oecd economic outlook norway, *Technical report*, OECD.

OECD: 2014, *Sharing the pain equally? Wageadjustments during the crisis and recovery*, OECD Publishing, chapter 2. url = "/content/chapter/empl$_outlook - 2014 - 5 - en$", .

Samuelson, P. A.: 1951, Economic theory and wages, *in* D. Wright (ed.), *The Impact of the Union*, Harcourt Brace & Company, pp. 312–43.

Solow, R. M.: 1990, *The Labour Market as a Social Institution*, Basil Blackwell, Oxford.

Soskice, D. and Iversen, T.: 2000, The nonneutrality of monetary policy with large price or wage setters, *Quarterly Journal of Economics* (1), 265–284.

Usher, D.: 2012, Bargaining unexplained, *Public Choice* **151**(1/2), 23–41.

Visser, J.: 2013, Wage bargaining institutions-from crisis to crisis, *European Economy. Economic Papers 488*, European Commision.