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DISCUSSION PAPERS

895

Discussion Papers No. 895, February 2019
Statistics Norway, Research Department

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Abstract:
Norway experienced a high immigration flow after the EEA directive in 2004 stating workers right to free movement within the European Union and EEA-countries. There is no clear consensus in the literature on how immigration affects native wages, but some studies using Norwegian micro data have estimated a negative effect of higher immigration for some type of workers. In this paper, to capture that the wage setting is highly coordinated in Norway, we model a system of native wages for three sectors; manufacturing, private service industries and public sector. We estimate that labour immigration has had a negative effect on the attainable wage growth for natives in all three sectors, but that the largest and most direct impact on wages has been in the private service industries. Immigration is found to be exogenous with respect to the parameters of our model of wage formation.

Keywords: Cointegration; Error-correcting adjustment; Estimation and hypothesis testing in cointegrated models; Macroeconomic fluctuations and transmission mechanisms; Short-run and long-run impact; Vector Autoregressive Processes; Pattern wage bargaining; Small open economy wage policies

JEL classification: C52, E24, E31 and J31

Acknowledgements: We would like to thank David F. Hendry, Ådne Cappelen, Jennifer Castle and Terje Skjerpen as well as seminar participants the Frisch sentre and Statistics Norway. This project has received financial support from the Norwegian Research Council (Grant number 236935). All numerical and graphical results in the paper were obtained by the use of OxMetrics 7.00, cf. Doornik and Hendry (2013a, b, 2014).

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ISSN 1892-753X (electronic)

Sammendrag

I løpet av de siste 10-15 årene har arbeidsinnvandring fra utenfor Skandinavia økt kraftig. Det er ingen konsensus om hvordan arbeidsinnvandring påvirker nordmenns (personer uten innvandringshistorikk) timelønn, se Borjas (2017), Dustmann et al. (2013) eller Nickell og Saleheen (2017).

I denne artikkelen har vi modellert nordmenns lønninger i tre sektorer; industri, private tjenester og offentlig sektor. Den økonomeriske modellen tillater oss å estimere mønsteret i lønnsdannelsen på nasjonalt nivå. Resultatene viser at det er et stabilt langsiktig forhold mellom lønnsnivået i de tre sektorene, med industrien som lønnsleder (såkalte frontfagsmodellen). Selv om dette er påvist tidligere med data fra nasjonalregnskapet, er det første gangen et slikt mønster estimeres med lønnsdata for nordmenn.

Estimeringsresultatene viser at lønnsnivået for nordmenn i industrien følger sektorens utvikling i pris og produktivitet og er mindre påvirket av innvandring enn det tidligere resultater som bruker nasjonalregnkapstall (Gjelsvik et al. (2015)). Denne forskjellen kan tolkes som at innvandring har påvirket industrilønnsgjennomsnittet ved å endre sysselsettingen. På den annen side finner vi at timelønningene er betydelig negativ påvirket av innvandring i privat sektor og i offentlige tjenester.

De estimerte effektene av innvandring synes ikke å være påvirket av endogenitetsskjevhet. Dette kan skyldes at den estimerte modellen er en delmodell, og at innvadringen er endogen i et annet, men tilknyttet system (Cappelen et al. (2015)). Dette større systemet må spesifiseres før en kan beregne mer presist effekten på lønnsnivåene av en varig økning av innvandringsstrømmen til Norge. I prakis betyr dette en emprisk makromodell, som f.eks. KVARTS. Imidlertid kan vi, med bakgrunn av resultatene i notatet, si at det langsiktige lønnsnivået trolig er redusert i privat tjenesteyting og offentlig sektor på grunn av skiftet i arbeidsinnvandringen. Innvandringen har også påvirket den relative lønnen mellom private og offentlige tjenester. Dette kan tyde på at innvandring har påvirket forhandlingsmakten (en parameter av systemet), men så langt uten å fundamentalt endre mønsteret i den nasjonale lønnsdannelsen.

1 Introduction

The years 2004 and 2007 mark two recent expansions of the European Economic Area (EEA), see Directive 38 (2004). In addition to Cyprus and Malta, eight Eastern and Central European countries joined the European Union in 2004. Bulgaria and Romania were added to the list of members in 2007. Differences in economic conditions between the old and the new member states gave rise to a clear migration pattern where the old member states were predominantly hosting countries. This allowed for a more efficient allocation of the labour force within the EEA. But this free movement was also met with concerns for increased labor market competition at the lower end of the wage distribution and increased competition for public goods (Kahanec et al. (2009)).

Empirical evidence of the effect of immigrants on the wage development on natives are mixed in Europe and the US, see Manacorda et al. (2012), Ottaviano and Peri (2012), Borjas (2017), Dustmann et al. (2013) and Nickell and Saleheen (2017). In a comment to the literature, Nickell and Saleheen (2017) write that empirical studies indicate that labour immigration may have affected wages of employed wage earners less than many commentators believed would happen. In a summary of their own investigation of the effect based on UK data, the same authors characterize the effects of immigration on wages as moderate, but negative, and that the results have an occupation and skill profile. However, results based on UK data, tell us little about which effects labour immigration may have had on a system of collective wage bargaining, simply because collective wage bargaining had been reduced and played a minor role in wage formation long before UK opened up to the European labour market. However, in several other European countries, national bargaining systems based on the agreements between the unions and employer organizations, are still a reality. Such systems can either adapt to the new realities of internationalized labour markets, or break down under the challenges of the changing working life, OECD (2018).

In the case of Norway, throughout the post-war period, bargaining in the manufacturing sector have defined the wage norm for the rest of the economy. It is the main organizations in manufacturing industry, that propose a wage norm ahead of the negotiations. It is expected that the following agreements within the manufacturing industry follow the wage norm, but there is no legal or other automatic extension mechanism that force them to stick to the wage norm. At the end of the day, the system is voluntarist, Evju (2014). It is reasonable to assume that the system can be disrupted if the gains from deviation are perceived by some of the partners to have increased.

Gjelsvik et al. (2015), looked at a multiple equations econometric model of the system of national wage formation in Norway and tested the robustness of the model equations with respect to labour immigration flow (gross labour immigration flow from EU/EFTA countries, North America, Australia and New Zealand and non-EU Eastern Europe to the population ages 15-74 in percent). The main conclusion was that while the pattern wage bargaining system were robust to the increase in immigration, the level of the wage norm in manufacturing sector was affected by the expansion in labour supply. The implication was however not that the slope of the long run wage-path was permanently affected. The growth rate of the long run wage norm was estimated to be the growth rate of the wage scope in the manufacturing sector. There was also an estimated effect of immigration in private service production.

However, the estimated effects of labour immigration flow on Norwegian National Accounts average hourly wage in manufacturing and the private sector may result from composition effects, i.e. if immigrants replace natives in lowest paid jobs average hourly wage is lowered. Hence, the previous results are difficult to interpret as evidence of reduced capability of collective action as a result of increased labour immigration during the first 15 years of the 2000s. In this paper, we revisit the modelling of the pattern wage bargaining system, but with data that have been specially constructed from micro data to represent the hourly wage for only native Norwegian workers.

The data set that we use is quarterly, from 1980(1) to 2014(4). This means that about 1/4 of the sample is from the recent period of increasing and high labour immigration. The Norwegian labour market was also influenced by immigration through an agreement of free labour mobility between the Scandinavian countries prior to 2004, though immigration flow was at a low level. Specifically, prior to 1997, where we lack micro data to construct hourly wage for natives, immigrants accounted for only 5 percent of total working hours, hence the composition bias in this period is considered to be small. After this, immigration increases rapidly and hence wage growth in National Accounts might have been be heavily influenced by the wage level of immigrants. We use the annual hourly wage growth of natives from register data from 1997 to 2014 to calculate average hourly wage growth of natives for the three sectors; manufacturing, public and private. The growth rates are used to prolong the hourly wage from National Accounts for each sector after 1997.

The paper is organised as follows. Section 2 discusses how we construct wage data for natives by using National Accounts and register data. Section 3 includes background information on wage formation that is important for the Norwegian wage setting. In Section 4 the data set and the long run hypothesis are presented. The results are in Section 5. The results indicate that the rise in immigration appears to have reduced the compromise wage level that the social partners in private service industries can coordinate on in a pattern wage bargaining model. In the concluding section, we venture an interpretation in terms of the system's capacity of adjustment, but within bounds defined by necessary critical levels of organization and reach of collective agreements.

2 Natives' wage data

We investigate if wages of native workers have been effected by the large inflow after the EU-directive. The number of immigrants may affect the labour market of present workers in several ways if immigrants and native workers are close substitutes in the labour market. One effect is that increased labour supply affects the wage setting process and reduce the bargaining power of natives, that in turn will decrease wages of natives (lower long run wage norm).

In the following, we use the gross immigration flow from the EU/EFTA countries, North America, Australia and New Zealand and non-EU Eastern Europe, in percent of the population aged 15-74, since the impact on native wages from this group is higher than from all immigrants, see Bratsberg et al. (2014). They find that the impact of immigration on wages depend on the inflow region of origin, and the effect is greater for workers from neighboring countries because these workers are closer substitutes for native workers. Micro (register) data illustrate the increase in the share of working hours provided by immigrants from these countries to the total number of working hours in three sectors of the Norwegian economy: the manufacturing (sector 1), private service production (sector 2) and public services (sector 3), see Figure 1. The largest increases are found in sectors 1 and 3, while the development has been more modest in sector 2.

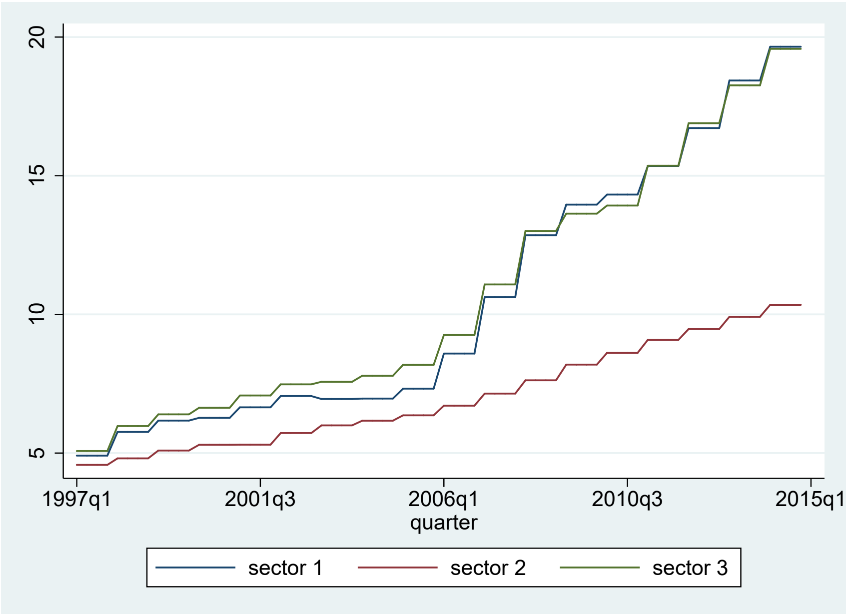

Figure 1: Working hours of immigrants as a share of total hours in manufacturing (sector 1), private services (sector 2) and public sector (sector 3)

The hourly wage as measured in the National Accounts (NA hereafter) may not be representative for natives if immigrants on average have lower hourly wages (composition effect). We use the micro (register) data information on country of origin to investigate the wage development of natives and combine this with wages in NA. It is unproblematic to combine the two data sources if they show the same development for the average wage for all workers in three sectors. Figure 2 shows time plots of hourly wage rates in NA relative to the hourly wage rates from register data adjusted for a break in the register data due to a change in data collection in 2009 (see Appendix for details). Figure 2 shows that there is a fairly stable ratio between the wages in the three sectors, although the ratio between the two sources in sector 1 has decreased from 1.3 to 1.25, and increased from 1.25 to 1.35 in sector 3 over the sample period.

We use the micro (register) data in order to construct time series data for the hourly wage level of native employees in three sectors of the Norwegian economy: the manufacturing (sector 1), private service production (sector 2) and public services (sector 3). The graphs in Figure 3 show the hourly wage for all workers in NA and the hourly wage for only natives (combination of NA and micro data). Figure 3, sector 1, shows that the hourly wage of natives in manufacturing are higher than hourly wages of all workers. The composition effect of immigrants is therefore negative in manufacturing industry. On the other hand, in public sector, hourly wages of natives are lower than the average of all workers, i.e. the composition effect of immigrants is positive. There is no difference between hourly wages of natives and the hourly average wage of all workers in private service industry, see Figure 3, sector 2.

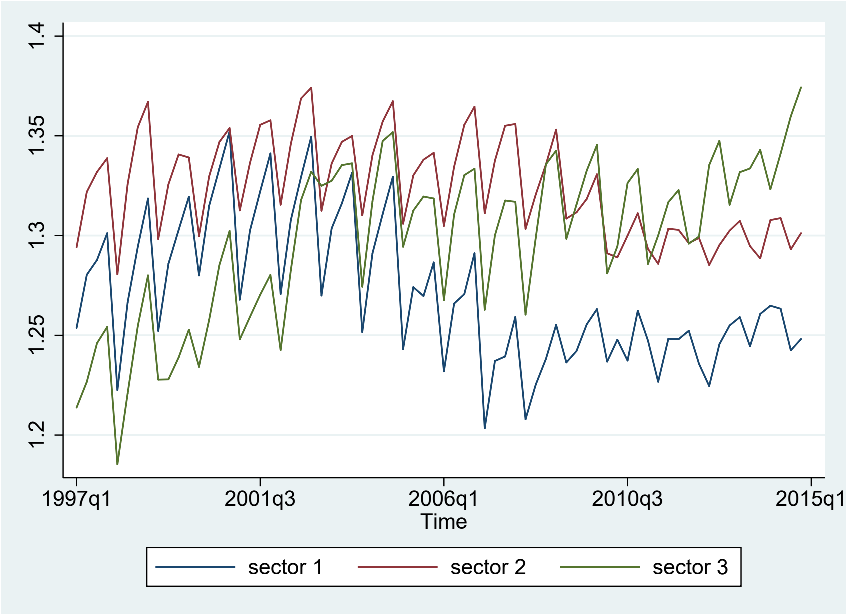

Figure 2: Hourly wage rates for all workers in NA relative to hourly wages in ATMLTO register in sectors 1, 2 and 3.

3 The national system of wage setting

In this section we give the institutional background of the collective bargaining system, which continues to be important for regulation of the wage level and the relative wages between broad groups of wage earners.

In Norway, around 66 per cent of all employees are covered by wage agreements in 2014, see NOU:10 (2017). The coverage of all employees in private sector is around 50 per cent, while 100 per cent of employees in public sector are covered.

The employees are organized in different trade union federations which are associated with four main organisations; The Norwegian Confederation of Trade Unions (LO), The Confederation of Vocational Unions (YS), The Confederation of Unions for Professionals, Norway (Unio), The Federation of Norwegian Professional Associations (Akademikerne) and about 20 smaller independent national organisations. LO was founded in 1899, and is by far the dominant union force. According to Løken et al. (2013); LO consists of 22 different national unions with a total of 895 257 members per 31 December 2012, YS consists of 20 independent trade unions with a total of 226 624 members, Confederation of 12 member unions with a total of 317 608 members and finally Akademikerne consists of 13 member organisations with a total of 170 387 members. LO has both skilled and unskilled workers as members in the same union and some unions organise both blue- and white-collar workers, but in parts of the private sector, blue- and white-collar workers are members of separate unions. Generally, each national union under LO covers a specific trade, branch of business or public service sector. Thus, the main organising principle is industrial unionism, but there are exceptions.

On the employers side there are five main organisations and some smaller independent organisations, where three of the main organisations are in the private sector. One of the three main private organisations, the confederation of Norwegian Enterprise (NHO), has 21 nationwide sector branch federations and 15 regional associations. The federations

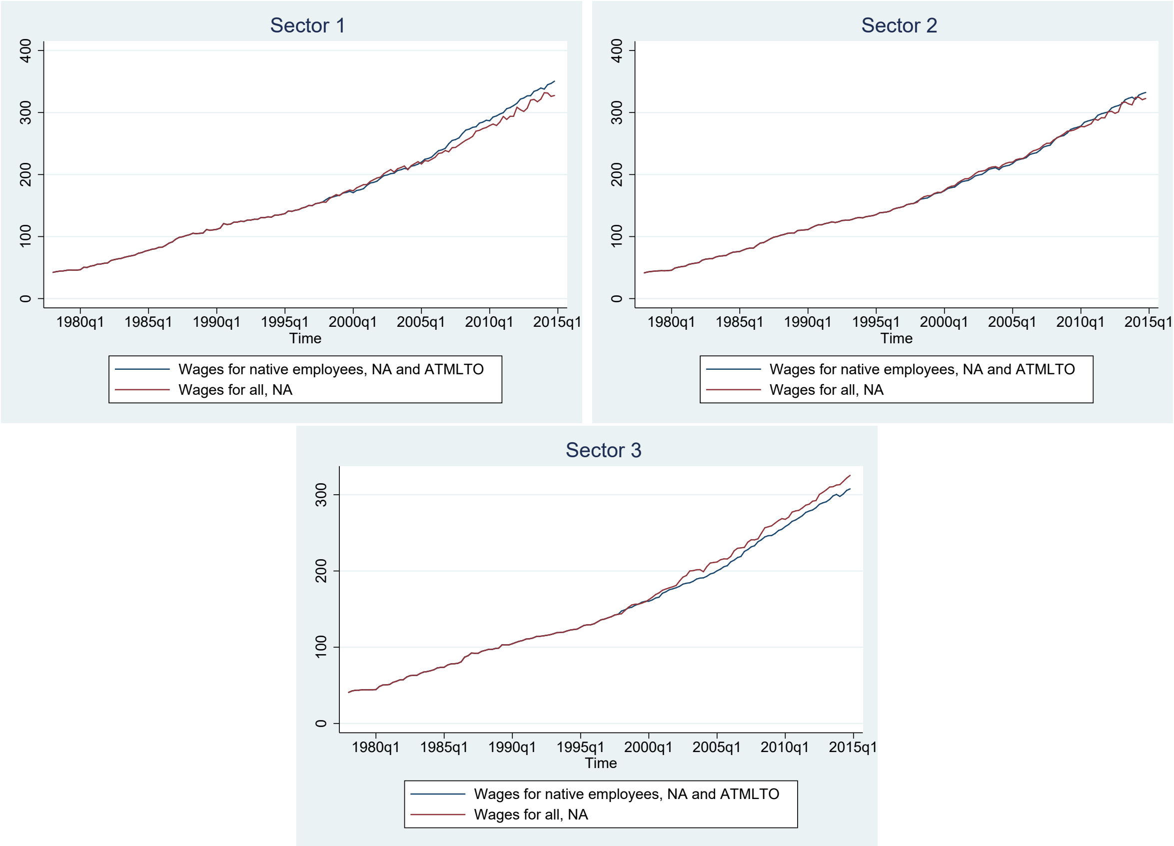

Figure 3: Manufacturing (sector 1), private services (sector 2) and public services (sector 3) wage formation, w_1, w_2 and w_3, and native wages for the same sectors, w_{N1}, w_{N2} and w_{N3}. Hourly wage rates are measured in NOK

negotiate separately with their counterparts, but NHO exerts a strong central authority over the federations regarding bargaining and the conclusion of collective agreements with LO and YS unions, and is party to all their collective agreements. In addition, many white-collar collective agreements are independent of branches and are renewed by NHO on behalf of several federations. The other two main private organisations bargain at a centralized and at a sector level with their counterparts.

The wage negotiations in Norway are therefore conducted at a lower level by the national organisations, but the wage setting is nevertheless judged as coordinated by experts, see Visser (2013, 2016). Throughout the post-war period, the settlements in the manufacturing sector have been perceived as those of the wage leading sector, and they are negotiated first. The manufacturing sector also signs the collective agreements in several other European countries, see Knell and Stiglbauer (2009).

In practice, it is the main organizations in manufacturing industry, the Confederation of Norwegian Business and Industry and the Norwegian Confederation of Trade Unions, that propose a wage norm ahead of the negotiations. As noted above, it is expected that organizations within the manufacturing sector follow the wage norm, but there is no legal sanctions or automatic extension mechanism in this voluntarist system. The wage negotiations are carried out by many different wage-setters and only about half of all employees in manufacturing industry (industrial workers) are settled before the following sectors start to negotiate. This implies that the wage followers do not know the actual outcome of the wage settlements in manufacturing sector before they decide on the wage growth. Again there are no sanctions or penalties for wage-setters in manufacturing or following sectors that decide to deviate from the wage norm. Thus, it is reasonable to assume that the system of coordination survives only as long as the parties in the wage settlement have something to gain from committing to the wage norm. The system can become disrupted, if the incentives change, or the ability to organize and coordinate is reduced.

4 Data set and the long-run hypotheses

The time series variables included in the empirical model is:[1]

$W_{it}-$Index for hourly wage for native workers 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, gross product 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 flow from EU/EFTA countries, North America, Australia and New Zealand and non-EU Eastern Europe, in percent of the population aged 15-74.
τ_t- represents the natural logarithm of the payroll tax rate plus one.

In the following, lower case (latin) letters refer to the logarithm of the original variables. For example, $u_t = \log(U_t)$ denotes the log of the unemployment rate.

Gjelsvik et al. (2015) provided evidence for unit root non-stationarity in the (logarithms of) hourly wage, price and productivity time series (in the NA version of the data set). Two other key time series, the (logarithms of) unemployment and immigration percentages are also non-stationary, but mainly as a results of breaks in mean, and not due to unit roots. The time plots of these two variables are shown in Figure 4. The graph illustrates a large inflow of workers after the EU-directive stating workers right to move freely within EU and EEA countries. However, also prior to 2004, workers from the Nordic countries could move freely between Norway and the other Nordic countries. As can be seen from the figure, a lower national unemployment rate is associated with higher immigration flow towards Norway, but the immigration percentage is shown to be weakly exogenous to the system that we formulate for wages, see section 5. In the following we keep the modelling assumption that the (logarithms of) hourly wages of natives are $I(1)$ variables.

A starting point when modelling the system of national wage setting is that the secular trend in wage costs per hour in manufacturing is defined by the value of labour productivity. The natural logarithm of the value of labour productivity, which we denote by x_t is by definition the sum of a price term, $q_t = ln(Q_t)$, and the average labour productivity in real terms, $z_{1t} = ln(Z_{1t})$:

$$x_{1t} \equiv \; z_{1t} + q_{1t}.$$

We assume that x_t is characterized by a unit root at the long-run frequency, meaning that the first difference Δx_{1t} is a stationary time series with time invariant mean. Following

[1] The data set is available from the authors on request.

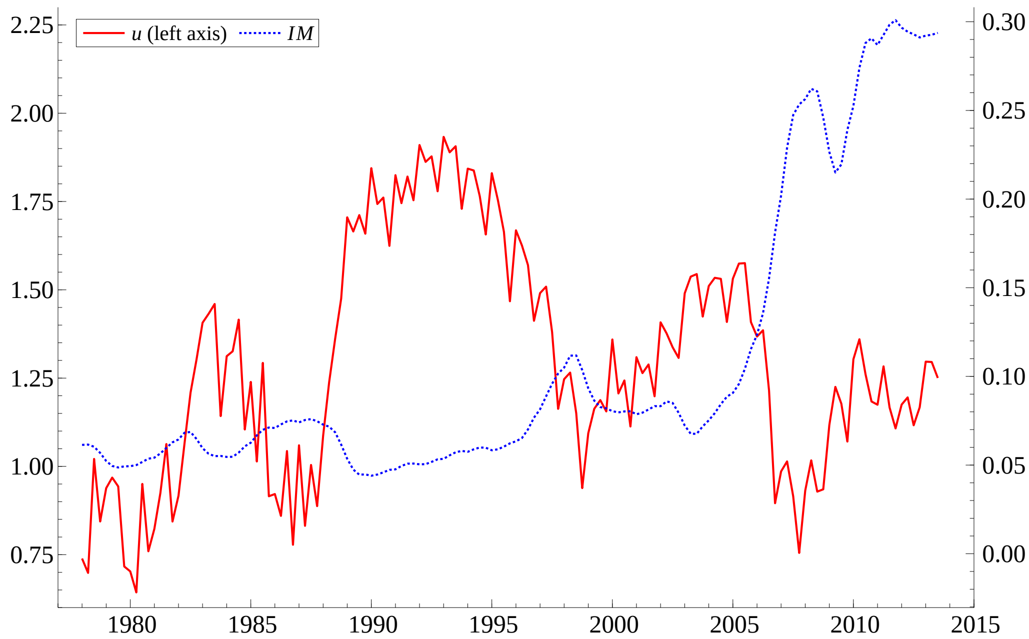

Figure 4: Plot of the natural logarithm of the unemployment rate (u), and of the immigration flow (IM) measured as gross immigration in percent of the population in working age

custom, we write this assumption as:

$$x_{1t} \sim I(1),$$

saying that x_{1t} is integrated of order one. The wage cost in manufacturing, which includes the payroll tax-rate τ_t, is defined as:

$$wc_{1t} =: w_{1t} + \tau_t.$$

Also this variable is assumed to be an $I(1)$ time series, $wc_{1t} \sim I(1)$, and hence $\tau_t \sim I(0)$.

The null-hypothesis is no relationship between wc_t and x_t, while the alternative hypothesis is that the two series are cointegrated:

$$H_0 : wc_{1t} - x_{1t} \sim I(1) \text{ , against: } H_1 : wc_{1t} - x_{1t} \sim I(0) \tag{1}$$

Economically, H_1 in (1) implies that, even if we can condition on productivity, we can only define a region for the real-wage in period t (*i.e.* a confidence interval). This implication is consistent with wage stickiness, which has become a standard assumption in modern macro economic theory, as in the Diamond-Mortensen-Pissarides (DMP) search matching model, see Hall (2005).[2] Although the terminology has changed, and the underlying theory has been developed, this also harks back to the earlier literature about a "wage-corridor", and that a sustainable path for the wage level in a small open economy is traced out by the trend in firms' ability to pay, and a "wage-scope" variable,

[2]Following Hall (2005), several papers have incorporated rigid wage setting in search models, see Gertler and Trigari (2009), Blanchard and Galí (2010) and Krogh (2016).

corresponding to x_{1t} in the notation introduced above, see e.g., Nymoen (2017), referring to Aukrust (1977). Informal evidence of a long-run relationship of this kind is provided by Figure 5, which shows a time plot of the manufacturing wage-cost series, together with the graphs for the two components of the wage-scope variable, *i.e.* q_{1t} and z_{1t}.

The expectation $E(wc_{1t} - x_{1t})$ is the long-run equilibrium wage-share in manufacturing. More generally, the equilibrium wage-share is likely to depend on several underlying factors, e.g., related to production technology, product market conditions and bargaining power. Bargaining power has been attempted to be modelled theoretically by the relative impatience of the parties (Rubinstein, 1982), but it unclear how well that concept matches the strategies adopted by the parties in the system of national wage setting, where compromises are more typical than conflicts.

Be that as it may, it is only when possibly several institutional and market factors are constant that $E(wc_{1t} - x_{1t})$ is likely to be a stable parameter. In line with this, several earlier studies over the last decades have also concluded that the log-run wage share (*i.e.* $E(wc_{1t} - x_{1t})$) can be modelled as a function of the logarithm of the rate of unemployment, and other variables that measure wide sense bargaining power and ability to compromise about nominal wage regulation, see e.g. Nymoen (1989a), Johansen (1995), Bårdsen and Nymoen (2003). In our study, following Gjelsvik et al. (2015), we include only one variable in addition to $u_t = ln(U_t)$, and that is the gross immigration flow in percent of the population aged 15-74, IM_t.

Hence, in equation form, the hypothesis that we use as the main alternative to no-cointegration for the manufacturing sector wage, is the long-run equilibrium relationship:

$$wc_{1t} - x_{1t} = \beta_{10} + \beta_{u1} u_t + \beta_{1IM} IM_t + e_{1t},\ \beta_{u1} \leq 0, \tag{2}$$

or equivalently:

$$w_{1t} = \beta_{10} - \tau_{1t} + x_{1t} + \beta_{u1} u_t + \beta_{1IM} IM_t + e_{1t}, \tag{3}$$

where $e_{1t} \sim I(0)$ under the (alternative) hypothesis of cointegration, while $e_{1t} \sim I(1)$ under the null-hypothesis.

We next turn to the hypothesized relationships between the manufacturing wage and the compensation of wage earners in other sectors of the macro economy, sometimes referred to as the wage-pattern. Under wage-coordination, the long-run relationships between the manufacturing wage and the two other wage rates in our data set can be represented by the relativities:

$$w_{2t} = \beta_{20} + w_{1t} + \beta_{2u} u_t + \beta_{2IM} IM_t + e_{2t}, \tag{4}$$

$$w_{3t} = \beta_{30} + w_{2t} + \beta_{3u} u_t + \beta_{3IM} IM_t + e_{3t}. \tag{5}$$

where $e_{it} \sim I(0)$ $(i = 2, 3)$ under the alternative hypothesis of cointegration.

Since wage formation is closely related to price formation in the medium run perspective, we also include a hypothesized cointegrating relationship:

$$p_t = \beta_{40} + \beta_{4w}(wc_{1t} - z_{1t}) + \beta_{4pi} pi_t + e_{4t}, \tag{6}$$

where $e_{4t} \sim I(0)$ under cointegration, but $e_{4t} \sim I(1)$ if cointegration fails.

Equations (2)-(6) are formulated as an ordinary simultaneous equations model (SEM), and the equations are identified on the standard order and rank conditions for SEM, see

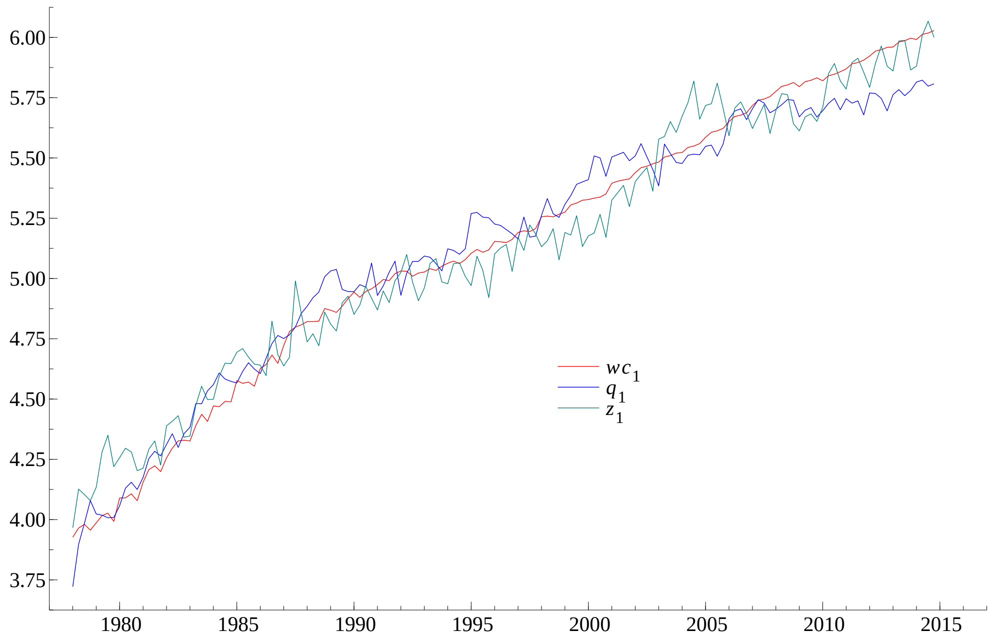

Figure 5: The manufacturing hourly wage cost index for natives $wc_1 \equiv w_1 + \tau_1$ graphed together with average labour productivity z_1 and value added deflator q_1. All three variables are measured in natural logarithms

Hsiao (1997). It may also appear to be a recursive system, as for example w_2, w_3 and p do not appear in the long-run equation (2) which is normalized on wc_1. However, assuming cointegration, the covariances are non-zero in general, and cointegration by its own implies that (Granger) causality must be present in the system, but does not imply "one way" causality. The direction of causality and the interpretation in terms of pattern wage bargaining was modelled and tested in detail by Gjelsvik et al. (2015) using National Accounts data. The results of that study were that both VAR-based tests and results from simultaneous equation modelling indicated that the manufacturing sector has acted as a wage leader both before and after the immigration influx.

In this paper we investigate however the implication of labour immigration on the long-run relationships of the system, when we use wage data for natives, as explained above.

5 Methodology and empirical results

We model wage per man hour for natives in three sectors (manufacturing, private services and public sector) and the domestic price level. The vector of modelled log-level variables is denoted by $\mathbf{Y_{wpt}} = (w_{1t}, w_{2t}, w_{3t}, p_t)'$, and we write the VARX (variously known as the open VAR) as:

$$\Delta \mathbf{Y_{wpt}} = \sum_{i=1}^{3} \mathbf{\Gamma}_i \Delta \mathbf{Y_{wpt-i}} + \mathbf{\Pi} \mathbf{Y_{t-1}} + \mathbf{\Upsilon} \mathbf{Q_t} + \boldsymbol{\varepsilon}_t, \qquad (7)$$

Table 1: Diagnostic test of VAR residuals, without deterministic trend, with p-values (in parentheses). Sample 1980(1)-2014(4)

	w_1	w_2	w_3	p
F_{AR}	0.45(0.81)	1.74(0.14)	0.53(0.75)	1.75(0.13)
F_{ARCH}	2.50(0.04)*	1.49(0.21)	14.61(0.00)**	3.43(0.01)
χ^2_{NORM}	2.37(0.31)	1.10(0.57)	7.10(0.03)	34.72(0.00)**

∗5% significance level.

∗∗1% significance level

where all variables in the long-run equations are in the data vector:
$\mathbf{Y_t} = (\mathbf{Y'_{wpt}}, x_t, \tau_{1t}, u_t, IM_t, z_{1t}, pi_t)'$, showing that we condition on six variables.

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 just noted, 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 driven by unit root non-stationarity.

$\mathbf{Q_t}$ contains other non-modelled variables in the VAR, indicator variables for income policy periods (Bowitz and Cappelen, 2001), reduction of the length of the working week (Nymoen, 1989b), constant term and three centered seasonals.[3] Following Johansen and Nielsen (2009), a main motivation for the use of indicator dummies is to robustify the estimators of the parameters of interest (in this case the eigenvalues associated with $\mathbf{\Pi}$). It has proven to be difficult to specify a VAR with residuals that are approximately normal in all respects, even with the conditioning and the dummies. Table 1 shows a representative test-battery where we include a constant and a deterministic trend. F_{AR} is the F-test for autoregressive residuals (order 1-5), see Harvey (1990, p. 174 and p. 278), Kiviet (1986); F_{ARCH} tests for ARCH of order 1-4, Edgren et al. (1969); χ^2_{NORM} is the test for departures from normality, Jarque and Bera (1980).

The diagnostic tests indicate that the VAR is relatively well-specified for the three wage growth variables, but that there is evidence of ARCH-effects and non-normality in the CPI-inflation residuals. Closer inspection shows that the non-normality of the residuals of inflation is due to kurtosis, and not skewness. It can be noted that the joint non-normality test for all four residuals (due to Doornik and Hansen (2008)) is insignificant at the 5% significance level.

The focus in the long-run analysis is the $\mathbf{\Pi}$ matrix with dimension 4×9. Let r denote the rank of $\mathbf{\Pi}$. Since x_t, u, IM, pi and z_1 are non-modelled variables, r can be 0, 1, 2, 3 or 4. $r = 0$ means that there is no cointegration, which would imply direct rejection of the long-run system we formulated above. If, on the other hand, $r = 4$, the variables x_t, u, IM, q_1 and z_1 represent five common trends in the three nominal wage rates and the price index. In this case, there is a possibility that the long-run equations above may represent an identified structure with empirical support.

The values of the trace test, which is the standard test for determination of cointe-

[3]The dummy variables in $\mathbf{Q}_t$ are: A dummy which is 1 from 1988(1) to 2002(4) and zero elsewhere. A dummy which is 1 in 1980(2)-1980(3),1981(2)-1981(4),1988(2)-1988(4),1990(2) and zero elsewhere. Indictor variables for 1990(3), 2011(1), 2012(1), 2013(1) and 2014(1).

Table 2: Tests of cointegration rank. Sample 1982(1)-2014(4)

Eigenvalue (λ_i)	H_0	H_1	Trace test statistics	5% Critical values
0.43	$r = 0$	$r > 0$	210.51	107.6
0.35	$r = 1$	$r > 1$	132.57	76.82
0.31	$r = 2$	$r > 2$	72.38	49.52
0.13	$r = 3$	$r > 3$	19.56	25.70

Endogenous variables: w_1,w_2,w_3,p.

Restricted variables: $x = (q_1 + z_1)$, u, IM, pi, z_1, trend.

Unrestricted: constant, τ_{1t}, seasonals and indicator variables.

5% critical values: Pesaran et al. (2000, Table 6(d)).

grating rank, are given in Table 2, together with 5% critical values from Pesaran et al. (2000, Table 6(d)). The values of the trace test statistic are well above the corresponding critical values for test of two or fewer cointegration relationships, which supports the assumption that the number of stationary long-run relationships is at least three. The last row shows that we are unable to formally reject three, and accept four, cointegrating vectors. Because of its role in the completion of an economically interpretable wage-price subsystem, we decided to proceed with four long-run relationships (i.e. $rank = 4$), although the fourth relationship has received weaker formal support than the three first.

In summary, the formal tests indicate at least three cointegration vectors. This result is central for us, as it gives formal empirical support for the idea of a system of long-run relationships between wages in three sector. We also consider that the relationship between wages, productivity, import prices and the CPI index is almost certainly part of wage-price process of a small open economy. Without putting too much weigh on the fourth relationship, the three first ones are the focus of our interest, we suggest to keep it as representation of domestic and foreign trends in CPI.

Using the decided cointegration rank, $\mathbf{\Pi}$ can be written as $\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where $\boldsymbol{\alpha}$ is $4 \times r$ and $\boldsymbol{\beta}$ $(9 \times r)$ is the matrix with cointegration parameters. The four relationships are however subject to further identification conditions before they can be interpreted economically, i.e., $\boldsymbol{\beta}$ is not identified from the trace test stage of the investigation, Hsiao (1997). However, the system of long-run equations (2)-(6) above is over-identified within the context of the VAR used for rank determinations. Therefore, in Table 3 we can present estimates of identified $\boldsymbol{\beta}_{id}$ matrices. Panel 1 in the table shows estimated versions of the three wage equations (2)-(5) and the price equation (6).

Below the estimated equations in Table 3, Panel 1, we report the log-likelihood value together with the Likehood-ratio (LR) test $\chi^2(15)$ of the 15 over-identifying restrictions. The LR-test is calculated against the benchmark likelihood reported as "Unrestricted Log L" in the "Memo" line at the top of the table (it is the maximised likelihood value after we have omitted the insignificant deterministic trend from the cointegrated VAR). As the unidentified (unrestricted) CVAR is heavily parameterized in this case, the rather low p-value of 0.06 of the LR test of the 15 restrictions is not surprising, and we take it as evidence that the main long-run structure receives empirical support.

Turning to the estimated coefficients of the gross immigration variable IM, Table 3, Panel 1 indicates that increased immigration reduced wages of native workers in all three sectors. However, judging by the t-values of the three estimated coefficients: -1.5,-3.2 and -0.9 the effect is only statistically significant in the private service sector. Panel 2 therefore reports the estimated long-run relationships after imposing the two additional zero restrictions. The joint test ($\chi^2(2)$) is statistically insignificant with a p-value of 0.6.

The estimated coefficient of IM in Panel 2 is numerically as well as statistically significant. Assuming a permanent change of 0.10 in IM, which is representative of what has taken place after 2005, the estimated resulting reduction of private service sector wages for natives is 6 %. If the relative wage in the public sector is to be maintained relative to the private sector, also the wage level in the central and local government administration has to be reduced by 6 %. Since 80 % of the workers belong to these sectors, the results indicate that the wage level for natives may have been broadly affected, even under the assumption that wage setting in the traditional wage-leading manufacturing sector has not been negatively affected by the opening up of the EU labour market.

Finally, we test the wage pattern system, and exclude the long run relationship of sectors 2 and 3 from the long run relationship for wages in sector 1 ($\alpha_{w1,ecm2} = \alpha_{w1,ecm3} = 0$), and similarly exclude the long run relationship of sector 3 from the long run relationship for wages in sector 2 ($\alpha_{w2,ecm3} = 0$). The joint test ($\chi^2(3)$) is statistically insignificant in Table 3, Panel 3.

Table 3: Estimated long-run relationships

Memo: Unrestricted $Log\ L = 2156.13877$

Panel 1: Identified long-run relationships

$$w_1 = \underset{(.)}{-1}\,\tau + \underset{(.)}{1}\,x \underset{(0.17)}{-0.92}\,u \underset{(0.98)}{-1.45}\,IM$$
$$w_2 = \underset{(.)}{1}\,w_1 \underset{(0.06)}{-0.37}\,u \underset{(0.32)}{-1.01}\,IM$$
$$w_3 = \underset{(.)}{1}\,w_2 \underset{(0.03)}{-0.14}\,u \underset{(0.15)}{-0.14}\,IM$$
$$p_t = \underset{(0.07)}{0.32}w_1 - \underset{(0.12)}{0.12}z_1 + \underset{(0.07)}{0.37}pi$$

$Log\ L = 2143.82571$, $\chi^2(15) = 24.626[0.06]$

Panel 2: Identified long-run relationships, $\beta_{w1,IM} = \beta_{w3,IM} = 0$

$$w_1 = \underset{(.)}{-1}\,\tau + \underset{(.)}{1}\,x \underset{(0.07)}{-0.31}\,u$$
$$w_2 = \underset{(.)}{1}\,w_1 \underset{(0.03)}{-0.21}\,u \underset{(0.09)}{-0.60}\,IM$$
$$w_3 = \underset{(.)}{1}\,w_2 \underset{(0.01)}{-0.06}\,u$$
$$p_t = \underset{(0.07)}{0.30}w_1 - \underset{(0.13)}{0.12}z_1 + \underset{(0.08)}{0.36}pi$$

$Log\ L = 2143.32024$, $\chi^2(17) = 25.637[0.08]$
Additional restrictions: $\chi^2(2) = 1.0109[0.6032]$

Panel 3: Identified long-run relationships and pattern wage bargaining,
$\beta_{w1,IM} = \beta_{w3,IM} = \alpha_{w1,ecm2} = \alpha_{w1,ecm3} = \alpha_{w2,ecm3} = 0$

$$w_1 = \underset{(.)}{-1}\,\tau + \underset{(.)}{1}\,x \underset{(0.06)}{-0.21}\,u$$
$$w_2 = \underset{(.)}{1}\,w_1 \underset{(0.02)}{-0.18}\,u \underset{(0.11)}{-0.61}\,IM$$
$$w_3 = \underset{(.)}{1}\,w_2 \underset{(0.01)}{-0.03}\,u$$
$$p_t = \underset{(0.07)}{0.36}w_1 - \underset{(0.13)}{0.17}z_1 + \underset{(0.08)}{0.33}pi$$

$Log\ L = 2141.06151$, $\chi^2(20) = 30.155[0.07]$
Additional restrictions: $\chi^2(3) = 4.5175[0.2107]$

P-values in square brackets.

As mentioned above, in a separate study (Gjelsvik et al., 2015) a similar wage-price system was estimated using National Accounts data. That study found significant effects of IM also in the long-run equation for manufacturing sector wages, but also mentioned that the effect might be wholly or partly due to employment composition effects. The above results support the interpretation that the previous paper's result for the manufacturing wage was mainly due to composition effects, not that the wage level of native's wage earnings in the manufacturing sector has become depressed by the labour immigration from Europe.

In Table 4, the estimated results of immigration and unemployment on native wages in Table 3, Panel 2 are compared to the results on hourly wage as estimated in Gjelsvik et al. (2015) using National Accounts data. We see that the largest difference between the two sets of results is for w_1. However, although the absolute value of the estimated coefficient of 1.67 for IM on National Accounts data was very large, it also had a wide confidence interval. Furthermore, compared to Panel 1 in Table 3 the results from the National Accounts data are not worlds apart. The main difference between the two studies is therefore that the wage data for natives make us more confident that the rate of employment remains a sufficient measure of labour market conditions for modelling of manufacturing wage setting. Taken together the cointegration analyses of the two data sets give empirical evidence that the wage level, and also the price level of the Norwegian economy, have become negatively affected by labour market immigration (by implication real wages have become less affected).

Table 4: Comparison of estimates using wage data for natives with estimates using National Accounts wage data. Sample 1980(1)-2014(4)

wages	National Accounts wage		Natives' wage	
(w_i)	IM	u	IM	u
w_1	-1.67^{*}	-0.66^{**}	.	-0.31^{**}
w_2	-0.21^{*}	-0.14^{**}	-0.60^{**}	-0.21^{**}
w_3	.	-0.01	.	-0.06^{**}

* 5 per cent significance level.

** 10 per cent significance level.

A recurrent issue raised in connection with studies like ours is that the estimated effects of immigration is biased because the immigration rate IM is endogenous. It goes without saying that the identification and estimation of a different statistical system, where IM is modelled as an endogenous variable, can give important new information about the relationship. However, the implications of "economy endogeneity" of IM, i.e. the classification of IM as exogenous or endogenous, for the focus parameters in this study is not clear. Specifically, IM_t may well be an exogenous variable with respect to the parameters of the cointegration vectors (i.e. normalized on the wages in each sector, w_1, w_2 and w_3) even if it is "economy endogenous". IM_t is exogenous in this sense if the lagged equilibrium correction variable

$$ECMw_{2t} = w_{2t} - w_{1t} + 0.21u_t + 0.6IM_t$$

is statistically insignificant in an empirical model equation for IM_t and similar for the other equilibrium correction variables for wages. The model equation below, DIM (first difference of IM) is estimated by OLS on the sample 1980(1)-2014(1) with seasonal dummies CS_i for each quarter i. The result shows that the estimated coefficient of $ECMw_{1t-1}$, $ECMw_{2t-1}$ and $ECMw_{3t-1}$ are estimated to be practically zero (and with low t-values in parantheses). The implication is that there is no empirical evidence for the view that the conditioning on IM distorts the estimation results for that variable in a long-run wage equation for native workers.

$$
\begin{aligned}
\text{DIM}_t \;=\; & \underset{(0.0499)}{1.21}\ \text{DIM}_{t-1} - \underset{(0.0519)}{0.5524}\ \text{DIM}_{t-2} + \underset{(0.000362)}{0.00284}\ \text{CS}_{t-1} \\
& - \underset{(0.00167)}{0.005485}\ \text{I:1998(3)}_t + \underset{(0.00171)}{0.006035}\ \text{I:2006(2)}_t + \underset{(0.00169)}{0.01003}\ \text{I:2006(3)}_t \\
& + \underset{(0.00174)}{0.009336}\ \text{I:2006(4)}_t + \underset{(0.00182)}{0.01386}\ \text{I:2007(2)}_t + \underset{(0.00181)}{0.006645}\ \text{I:2008(1)}_t \\
& - \underset{(0.00168)}{0.005965}\ \text{I:2008(3)}_t - \underset{(0.00171)}{0.009078}\ \text{I:2008(4)}_t + \underset{(0.0018)}{0.008313}\ \text{I:2009(3)}_t \\
& + \underset{(0.00185)}{0.00831}\ \text{I:2009(4)}_t - \underset{(0.00184)}{0.006904}\ \text{I:2010(1)}_t + \underset{(0.00177)}{0.01349}\ \text{I:2010(2)}_t \\
& + \underset{(0.00171)}{0.009361}\ \text{I:2011(2)}_t - \underset{(0.000117)}{2.874e-005}\ \text{ECMw}_{1t-1} + \underset{(0.00236)}{0.0009101}\ \text{ECMw}_{2t-1} \\
& + \underset{(0.0066)}{0.004763}\ \text{ECMw}_{3t-1}
\end{aligned}
$$

Estimation of this equation on a sample that ends in 2003(4), gives the same kind of results: the coefficients of $ECMw_{1t-1}$, $ECMw_{2t-1}$ and $ECMw_{3t-1}$ are $0, 0.0016$ and 0.0028 with t-values below 0.7.

Clearly, since the wave of European labour immigrants only arrived late in our sample, after 2005, these estimates must be interpreted with care, as a first generation of estimates of labour immigration effects on aggregate wages. However, the recursive estimates of the coefficients in Panel 2 of Table 3 show that the effect of immigration stabilises when immigration increases, see Figure 6.

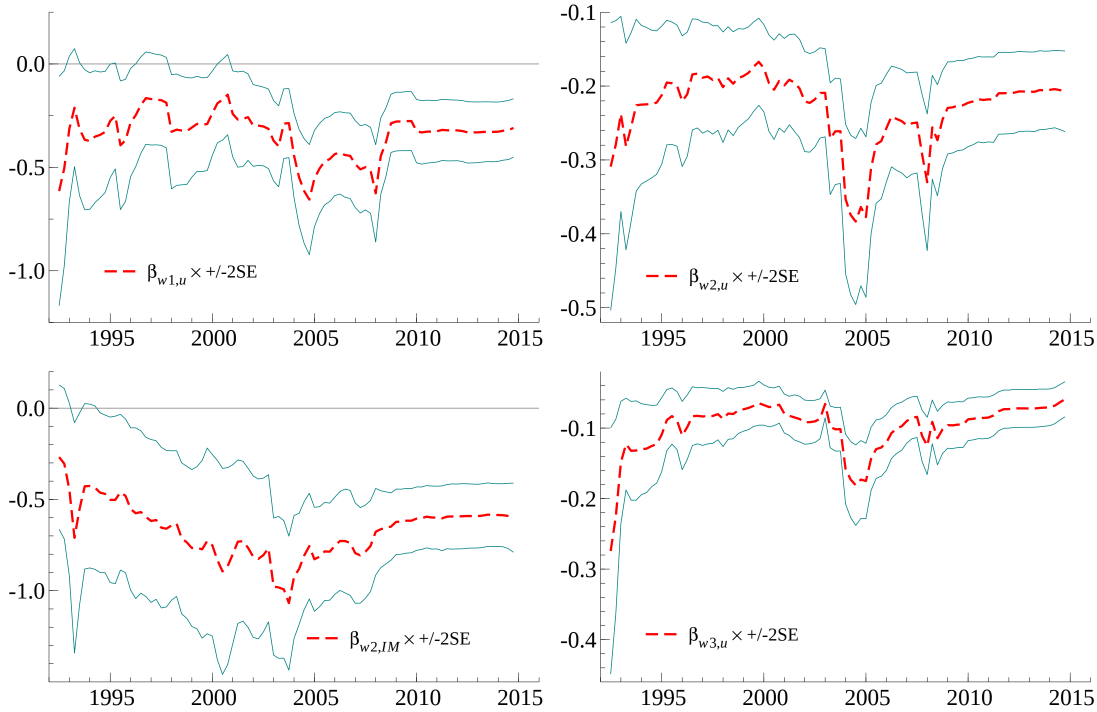

Figure 6: Recursive estimates of the long run coefficients of unemployment and immigration on wages in sector 1, 2 and 3

6 Conclusion

In this paper we have modelled wages of natives in three sectors of the Norwegian economy; manufacturing, private services and public sector. The model represents the typical pattern of wage bargaining in Europe, where manufacturing sector negotiates first and where the outcome is considered to lead the outcome of the other sectors of the economy. The logic of this pattern is simple, if the wage growth in the exposed sector leads the wage setting in the other sectors, the functional income distribution will be relatively stable both in the wage-leading and wage-following sectors.

Over the last 10-15 years, the rise in labour immigration from outside Scandinavia has increased sharply. There are no consensus in how labour immigration affects wages of natives, see Borjas (2017), Dustmann et al. (2013) or Nickell and Saleheen (2017).

The empirical results in this paper show that the wage growth in manufacturing native wages depend heavily on the profitability growth in the same sector, which is consistent with the theory of collective wage bargaining. Compared to an earlier study, Gjelsvik et al. (2015), the wage level in manufacturing has been less affected by immigration than what the results that use National Account data indicate. We interpret this difference as being due to employment composition effects within the sector. On the other hand, we find that wages of natives are significantly negative affected by immigration in the private service sector and in the public services. Labour immigration as a share of population 15-74 has increased from 0.1 to 0.3 percent after 2004. This has reduced wages in private service industry with 12 percentage points.

A recurrent issue is that the estimated effects of immigration may be biased because the immigration rate IM is endogenous. We find no empirical evidence of immigration being endogenous to the system, but the long term effect may be uncertain if the level of immigration depend on the development of the Norwegian economy, see Cappelen et al. (2015).

In addition, the lower wage level for some sectors may change the coordination of wage setting. The mechanism of the wage pattern bargaining system in Norway relies on that the many wage setters gain more in the coordinated solution than in the non-coordinated solution. Reduced gain may therefore alter the voluntary coordination of wage setting and/or 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 cointegration analysis showed that there is a stable long term relationship between wage levels in the three sectors, with the manufacturing sector as the wage leader.

According to the results, the long term (wage norm) wage level became lowered in private service industry and public sector due to the shift in labour supply trough immigration. In our model, that can change back again (if the other characteristics in the system are intact). 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.

A Use of register data

Register data contain information on employee country background and hence wages in NA can be decomposed into two groups for natives and immigrants by using register data from 1997 to 2014 given that the average wage growth rate in total is in accordance with the growth figures from NA, see Figure 7. However, the expected number of hours worked was registered for three categories of workers (between 4 and 19.9 hours per week, between 20 and 29.9 hours per week, and at least 30 hours per week) in ATMLTO before 2009 while expected working hours measured on a continuous scale are reliable from 2009 onwards. Wages in NA is approximately 0.3 percentage points above the measured wage in ATMLTO before 2009 and 1.2 after 2009. The drop is somewhat larger for sector 2 than the two other sectors. We have therefore adjusted the figures in the three sectors to avoid the large change caused by the change in data collection, see Figure 2.

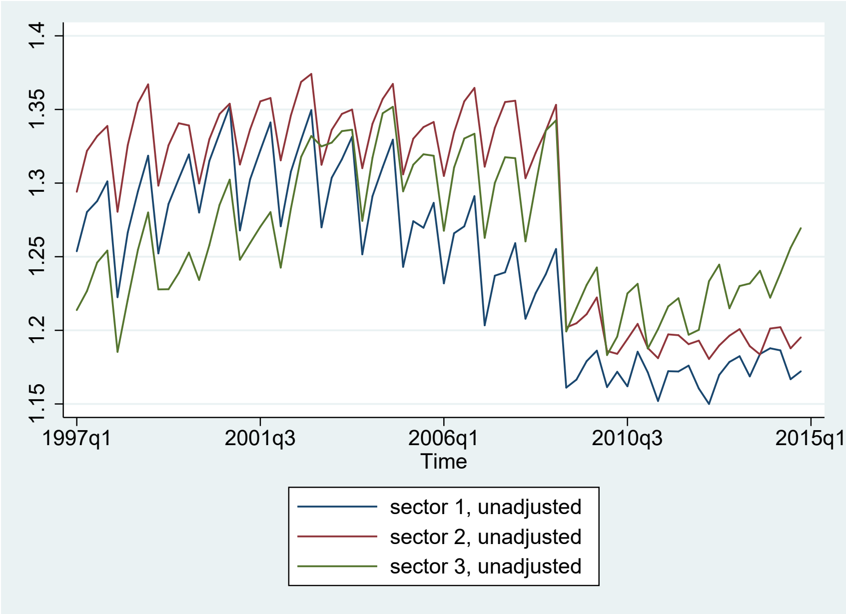

Figure 7: Wages in NA relative to wages in ATMLTO register in sectors 1, 2 and 3

In doing this we have prioritized not to distort the information that the register data (ATMLTO) have about the growth rates for the wages of natives. We have therefore regressed logged wage rates on a constant, a step dummy after 2008 and year dummies, and corrected the level according the value of the step dummy.

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