UiT The Arctic University of Norway

School of Business and Economics

# Cointegration in the Norwegian Model for Wage Formation

Against the Backdrop of the 2004 EU Expansion and the 2001 Change to Inflation Targeting, Have Wage Levels in the Construction and Wholesale Undersectors Continued to Exhibit a Long-Run Relationship with Wage Levels in the Industry, Public and Private Sectors?

Kristoffer Viland

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## PREFACE

This thesis does not only mark the end of a two-year master's program in Economics at the University of Tromsø - Norges arktiske universitet, but also seven years as a student at the university. Each of the seven years has added to the shaping of me academically and personally.

I would like to thank my supervisors Øystein Myrland and Espen Sirnes, Professor and Associate Professor at the Business School of UiT, respectively, for their support and useful comments throughout the writing of this thesis. The topic of the thesis, as well as the data, have been provided by Ådne Cappelen, researcher at Statistics Norway. Hence, I would like to extend a special thanks to Ådne for his invaluable insights and our productive discussions.

**Abstract**

In this thesis, hourly wage rate data spanning 1978-2018 has been collected to investigate if the wage levels in the construction and wholesale sectors follow the wage levels in the industry, public, and private sectors. Against the backdrop of the EU expansion in 2004 and the 2001 implementation of inflation targeting monetary policy, there has been an interest in studying the stability of the dynamics of the Norwegian wage formation model. Anchored in data, the construction and wholesale undersectors have been selected to study the cointegrating dynamics of the Norwegian model. Testing for cointegration in the variables, the results show one cointegrating vector; implying the existence of a long-run equilibrium relationship between the wage variables. The results thus show that, despite the EU expansion and the introduction of inflation targeting monetary policy, the Norwegian model ensures that the sector-wise wage levels never stray too far away from each other. The econometric tests employed in this thesis have been performed using the `urca` package (Pfaff, 2008), in the statistical software `R`. The figures have been created in `Python`, using the package `matplotlib` (Hunter, 2007). The tables have been created in the typesetting system LaTeX, using the package `booktabs` (Fear, 1995). The thesis in its entirety has been written and compiled in `R markdown`.

*Keywords:* wage formation, main-course model, cointegration, inflation targeting, labour immigration

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# 1 Introduction

Wage formation policy in Norway has, in large part, been governed by the so-called 'Frontfagsmodellen', or, the 'Main-Course Model', in which wage negotiations are first conducted in the industries primarily exposed to foreign competition, and the resulting wage settlement acts as a wage norm for the overall economy. In this manner, the profitability of the exposed industries is maintained compared to foreign competition. These factors, combined with the requirement of estimated equal return on capital to international trading partners, lead to the conclusion that Norwegian real wage growth cannot surpass the productivity growth in the exposed sector (NHO, 2017).

Two structural changes entered into the Norwegian economy in 2001 and 2004, when inflation targeting was implemented, and the enlargement of the European Union (EU) was ratified, respectively. Inflation targeting was a notable change in Norway's monetary policy, and the EU enlargement caused a surge of immigration into Norway, a considerable fraction of which was labour immigration. Naturally, this led to a significant labour supply shock in the economy. These structural changes introduced challenges to the already well-established main-course model, which had already proven to be robust through setbacks in the Norwegian economy. Gjelsvik, Nymoen, and Sparrman considered these structural changes in a 2015 research paper, where they investigated the degree of invariance of these changes on wage formation in the industry, public and private sectors. Specifically, they found that the dynamics of the model still constitute a long-term relationship between wage levels in the three sectors. However, the long-term wage level in the industry sector, as well as the relative wage between the public and private sectors, were negatively impacted by the high immigration levels. Furthermore, inflation targeting did not significantly affect the main-course model, neither did inflation expectations carry more importance in the new system.

The literature on the effects of the structural changes on the dynamics of the main-course model is limited. As such, there is an underlying interest in investigating the effects on more sectors of the economy. In this thesis, the literature will extend to include the construction and wholesale undersectors. Specifically, based on the results of Gjelsvik *et al.* (2015), the investigation is extended to test if the wage levels in the construction and wholesale undersectors have followed the wage levels in the industry, public and private sectors. The inclusion of these undersectors is relevant to the expansion of the literature on the current performance of the Norwegian wage model. Moreover, by introducing specific sectors to the investigation, we might obtain a more extensive foundation regarding the performance of the Norwegian model under various shocks and structural economic changes.

Chapter 2 will comprise the background of the thesis, where emphasis will be put on the two structural changes. In chapter 3, the necessary theory surrounding the topic will be explored; in particular cointegration and nonstationary variables. Subsequently, chapter 4 will disclose the variables and the methodology of applying the theory and models. The results will be presented and discussed in chapter 5, and finally, concluding remarks will be detailed in chapter 6.

# 2 Background

The contents of this thesis will largely be based on the findings and conclusions disclosed in the 2015 paper *Have inflation targeting and EU labour immigration changed the system of wage formation in Norway* by Gjelsvik, Nymoen, and Sparrman. As such, in some respects, this thesis can be regarded as a spin-off paper of the 2015 paper, expanding on the cointegration aspect by including more sectors in the analysis.

In the introductory section, the two structural changes; inflation targeting monetary policy in 2001 and EU's expansion in 2004, were described. In Figure 1, these have been marked in a plot with the evolution of the hourly wage rate of the industry, public and private sectors. The two events will be discussed in section 2.2.1 and 2.2.2. As already stated in the introduction, this thesis will investigate if the hourly wage rates in the construction and wholesale undersectors have followed these three sectors. However, a justification for the addition of the undersectors should be included. In 2012, Bratsberg and Raaum published the paper *Immigration and Wages: Evidence from Construction*, in which they used panel data to investigate the wage impacts of immigration in the construction sector. They find that a 10% increase in immigrant employment reduced the wages of Norwegian workers by 0.6%. Bratsberg and Raaum continued their studies on immigration impacts on wages in 2013 when they published an article investigating the marked increase in Swedish labour immigrants from 1990-2010. Among their conclusions, they find that the wholesale sector was one of the most supplied workplaces by the Swedish immigrants. While the surge of Swedish immigrants was not directly caused by the EU expansion, it is still of relevance in regards to the inclusion of the wholesale sector in this thesis. In a 2019 report from the Norwegian Confederation of Trade Unions (LO), the wholesale sector is noted as one of the sectors most affected by an increase in foreign labour. Particularly, in the fraction of the population aged 20-66, the fraction of employed labour immigrants increased by 56% from 2004-2014 in the wholesale sector. Similarly, this percentage was 75% in the construction sector. In the period 2015-2018, the percentage of employed labour immigrants in the wholesale sector increased by 13%, whereas it increased by 25% in the construction sector. See Tables 1 and 2 for more detailed information.

TABLE 1: Change in Employment. Q4. 2004-2014 (20-66 years)

| | Immigrants | Population excl. immigrants |
|---|---|---|
| Construction | +75% | +2% |
| Wholesale | +56% | −4% |
| Transport | +55% | −11% |
| Accommodation and serving | +70% | −6% |
| All sectors | +51% | −0.5% |

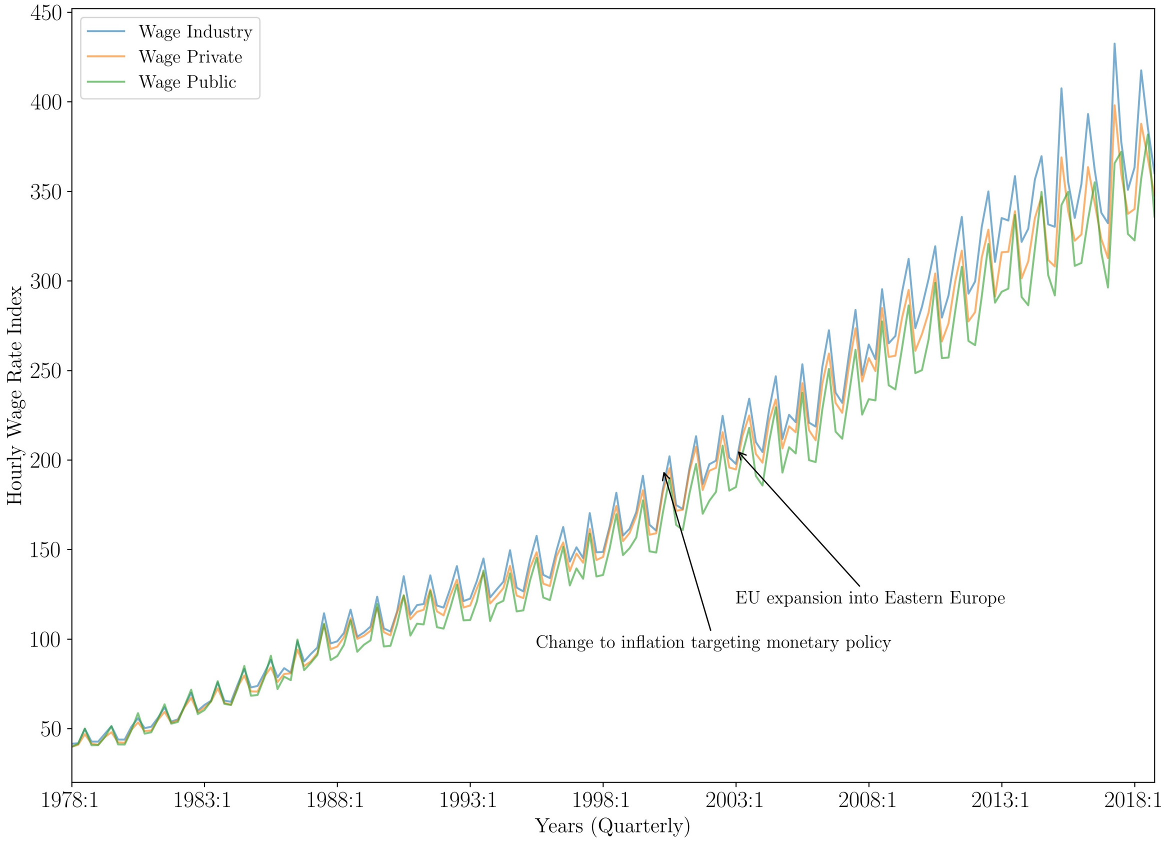
*Source:* NHO (2019, p.12) *Et endret arbeidsmarked - Hva betyr innvandringen?*

TABLE 2: Change in Employment. Q4. 2015-2018 (20-66 years)

| | Immigrants | Population excl. immigrants |
|---|---|---|
| Construction | +25% | +6% |
| Wholesale | +13% | −0.3% |
| Transport | +12% | −6% |
| Accommodation and serving | +13% | +4% |
| All sectors | +14% | −1.5% |

*Source:* NHO (2019, p.12) *Et endret arbeidsmarked - Hva betyr innvandringen?*

FIGURE 1: Structural Changes



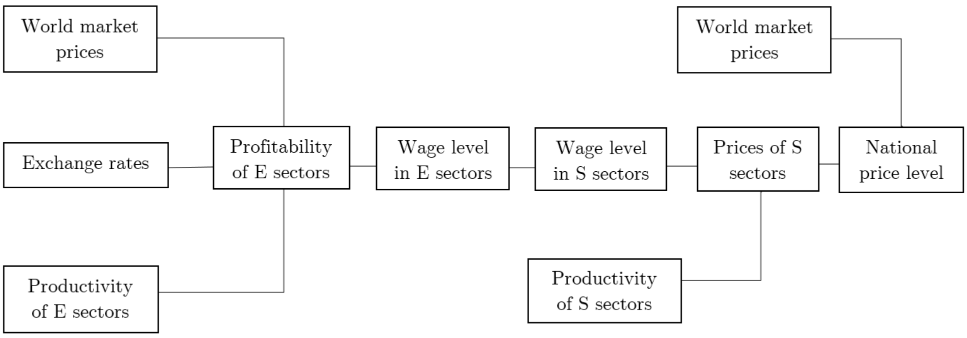

*Source:* The data has been collected from the KVARTS model variable list, provided by SSB.

## 2.1 Wage Setting in Norway

The notion of Wage setting in Norway springs from the ideas and research disclosed by Statistics Norway (SSB)—then the Central Bureau of Statistics in Norway—throughout the 1960s. Chief among these is the fundamental distinction between the *exposed* and *sheltered* industries, first detailed in SSB's *Economic Survey 1962*. Briefly, these industries are respectively defined as "*...the industries that are exposed to strong competition from abroad, either because they export most of their products or because they sell their products on the domestic market under strong foreign competition*", and "*...those whose products are marketed at home under conditions that leave them relatively free from foreign competition.*" (Aukrust, 1977, p.109-110).

The distinction is sound, as analyses of prices and incomes in these exposed and sheltered sectors are different in nature; whereas it is the global market which regulates the output prices of the exposed sectors, therefore making it impossible to compensate for increased costs through price adjustments, the sheltered sectors need not worry about foreign competition and will adjust prices in response to cost increases (Aukrust, 1977). See Figure 2 for a visualization of the dynamics in such a two-sector model. Exposed sectors are denoted by “E” and sheltered are denoted by “S”.

FIGURE 2: Illustration of the Mechanisms of the Two-Sector Model




*Source:* The figure is a reconstruction of the figure ’two-sector, long-run model mechanisms’ first instroduced in Aukrust (1977).

The figure is a recreation of the figure first sketched in Aukrust (1977), p.113.[1] In the figure, dynamics of price determination on the national level and the mechanisms involved in determining wage levels in the exposed and sheltered sectors are presented. The regulative mechanisms ensure that the wage levels in the exposed sector do not hurt its competitive ability in the international market.

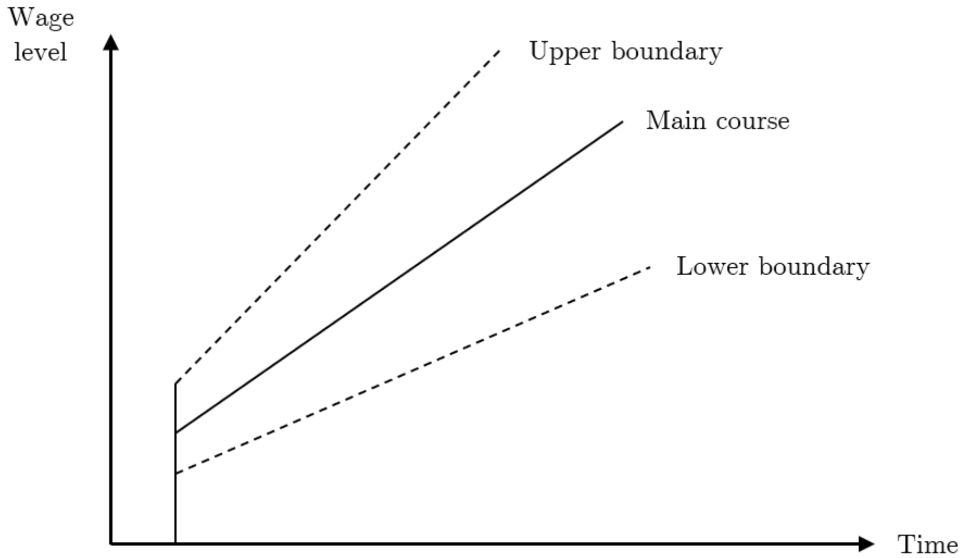
Aukrust (1977) notes the deviating nature of the profits in the exposed sector, and how these are corrected by different forces: wage negotiations, as profit levels are important aspects when negotiating wages are concerned; market forces, which also corrects wages to normal levels (illustrated by imagining very low wage costs, which lead to abnormally high profits, leading to excess labour demand causing a wage spike which again corrects wages back to normal levels); and economic policy, as the competitiveness of the exposed sector must be maintained. These correcting mechanisms led Aukrust to coin the term “the wage corridor”. An illustration is presented in Figure 3.

The wage levels are limited by an upper and lower bound. Thus, when the wage levels are nearing the bounds, the previously explained correcting mechanisms will work to force the wage levels back normal levels.[2]

[1] A step-by-step walkthrough of the dynamics of the figure is also detailed on p.113-114

[2] The figure is presented in Bårdsen *et al.* (2005), but is a simplified version of the figure sketched in Aukrust (1977)

FIGURE 3: The 'Wage Corridor'




*Source:* The figure is a reconstruction of the 'wage corridor' figure first introduced in Aukrust (1977).

## 2.2 The Norwegian Main-Course Model

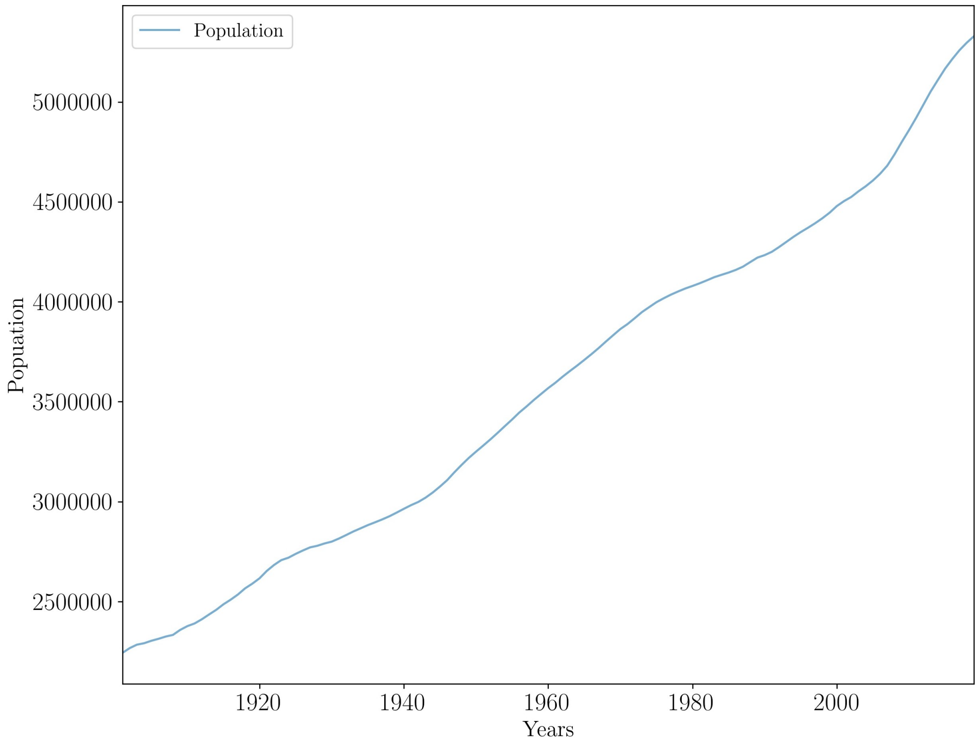
Section 2.1 largely establishes the framework of the Norwegian model of wage-setting. The essence of the Norwegian model is that the wage growth in the internationally competitive sector will act as a wage norm for wage negotiations in the other sectors. Since its inception, the model has survived several critical events that could have threatened its dynamics, including the international stagflation of the 1970s, and the Norwegian house- and banking crisis in 1989-1990 (Gjelsvik *et al.*, 2015). The model's robustness is cemented in the 2013 Official Norwegian Report *Wage formation and challenges for the Norwegian economy*, written by a select group of experts, where, despite facing structural changes and new economic development, the model's core still stands. In sections 2.2.1 and 2.2.2, the structural changes imposed in 2001 and 2004 will be further detailed.

### 2.2.1 Increased Labour Immigration

The population growth in Norway has been steadily increasing since the start of the 20th century. This growth has been plotted in Figure 4. From visually inspecting the graph of the development of the Norwegian population growth from 1900 to today, we can note some events and time periods that were particularly impactful. We saw the population growth halt from around 1920 to 1940, likely due to the interwar period; it was characterized by international uneasiness and an unstable peace treaty. Additionally, Norway was not able to share in the economic growth of the 1920s, likely due to an unfortunate deflationary monetary policy (Gjerde, 2018).

As with many other countries after World War II, the birth rate in Norway soared. Despite the “peace effect” being a temporary one, 1946, as noted by Østby (1995), remains the year with the highest rate of births to date. After the so-called “baby boom” period, the birth rate slacked off, and in 1983 Norway had the reported lowest fertility rate yet (Østby, 1995).

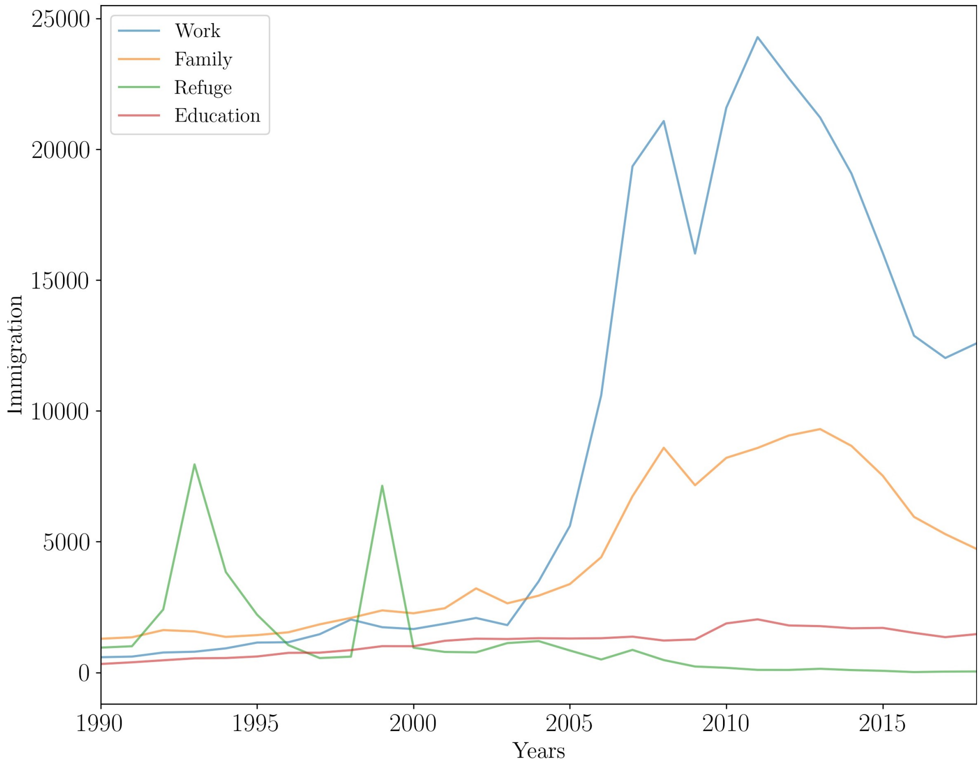
FIGURE 4: Population Growth in Norway




*Source:* The data has been collected from SSB—StatBank Norway.

The demand for work overtook the mainland supply of workers, and from 1970 immigrants increasingly covered this demand. The population continued to grow after Norway ratified the European Economic Area Agreement in 1994. And so, following the EU-enlargement into Eastern Europe in 2004, the population growth in Norway again soared. The activity in the Norwegian economy increased, which was reflected in the high demand for labour; Figure 5 graphs the different reasons for immigration to Norway. Consequently, a significant part of labour immigration into Norway went to the sector with the highest demand for labour—namely the construction sector (Holden *et al.*, 2013).

FIGURE 5: Reasons for Immigration




*Source:* The data has been collected from SSB—StatBank Norway.

### 2.2.2 Inflation Targeting Monetary Policy

In 1946, Norway joined the Bretton-Woods system of fixed exchange rates and employed this system until its collapse in 1971, upon which the Norwegian krone was allowed to float. This change was only temporary, however, as Norway signed the newly IMF-established system of fixed exchange rates, the Smithsonian Agreement, in 1972. From 1972-1978 Norway signed into the so-called *snake in the tunnel* monetary cooperation, where the currencies of countries within Europe were allowed to fluctuate to half of the bandwidth agreed upon in the Smithsonian agreement. Norway left the *snake* cooperation in 1978 and linked the krone to a trade-weighted currency basket. In 1990, the Norwegian krone was fixed to the European currency unit but was allowed to float after 1992. Finally, in 2001, inflation targeting was introduced with a target of 2.5 percent. (Alstadheim, 2016).

Soikelli (2002) notes that countries implementing inflation targeting in the early 1990s had previously struggled with either high and volatile inflation or high pressure on the exchange rate. After establishing an inflation target, many of the countries experienced a decline in inflation and good growth, creating political support for these regimes. Norway, however, which shared many of the characteristics of earlier implementors of the inflation targeting regime: a small and open economy; strong commodity share in exports; a centralized wage bargaining framework, implemented the regime on the tail-end of a strong boom in the economy. The labour force participation was high and inflation had been maintained at a low level, however, further economic growth seemed difficult without increasing pressure on the inflation. Given higher international inflation and low pressure on the krone combined with a strong central bank, the Norwegian transition to inflation targeting was uneventful (Soikelli, 2002).

In Inflation Report 3/2002, Norges Bank voiced their concerns on the potentially disruptive effects inflation targeting monetary policy would have on the current model of wage formation in Norway. Their concerns included the fact that under the fixed exchange rate monetary policy, it was the wage growth which would form the equilibrium point to maintaining international competitiveness in the Norwegian economy, whereas with an inflation target and floating exchange rate policy, it is the inflation target which determines the equilibrium level of wage growth over time. However, in their paper on the effects of inflation targeting and labour immigration on the Norwegian system of wage formation, Gjelsvik *et al.* (2015) empirically report that inflation targeting has had a limited impact on the structure of wage formation. Furthermore, the inception of the Technical Calculation Committee for Wage Settlements (TBU) in 1967 supports this, as inflation expectations have been an important factor in their operations.

## 2.3 Econometric Models of Macroeconomic Relations

The notion of using econometrics to model macroeconomic models and concepts can be studied in the 2005 book by Bårdsen *et al.*, among others. In chapter 3 in the aforementioned book, Aukrust's main-course model is introduced econometrically. The reconstruction of the theory is extensive and replication of this will not be necessary for the context of the thesis question. It is sufficient to state that the full econometric reconstruction of the Norwegian model can be studied in the book by Bårdsen *et al.*.[3]

What is essential in regards to this thesis, is long-run stable relationships between nonstationary variables, also known as cointegration. The process of regressing with non-

[3] In this thesis, the use of the wordings "the Norwegian model" and "the main-course model", are used interchangeably.

stationary variables, *spurious* regression, was discussed in Granger & Newbold (1974). Preceding their findings, regressions of this form were a relatively common occurrence in papers, and equations with such nonsense symptoms were wrongly presented with some worth. Granger continued his work on this topic and in 1981 published the paper *Some Properties of Time Series Data and Their Use in Econometric Model Specification*, where he suggested the concept of cointegration. However, cointegration was not formally introduced until a few years later, in 1987, when Engle and Granger published the paper *Co-Integration and Error Correction: Representation, Estimation, and Testing.* Using empirical data, they find, among other results, cointegration between consumption and income, but no cointegration between wages and prices. Johansen (1988, 1991) furthers the literature by introducing a maximum likelihood ratio test for a multivariate vector autoregressive process; allowing the presence of more than one cointegrating vector. Furthermore, akin to cointegration, the concept of *error-correction* was also suggested in Granger (1981) and formally introduced in Engle and Granger (1987). When there is a cointegrating relationship between two or more nonstationary variables, one can always estimate an accompanying error-correction model. The dynamic of the error-correction itself is straightforward to justify, as if two variables grows or moves in a similar manner, it is easy to imagine that there would exist both short-run adjustments and long-run adjustments that helps prevent the variables straying too far away from each other.

In 1989, Ragnar Nymoen published the paper *Modelling Wages in the Small Open Economy: An Error-Correction Model of Norwegian Manufacturing Wages*, in which cointegration theory is applied, and a dynamic wage equation in error-correction form is formulated. The paper investigates the manufacturing sector and finds that there are significant short-run effects from such variables as consumer price growth and normal work hours. In this thesis, the cointegration analysis is broadened and more general, as the analysis includes single wage rates for different sectors.

# 3 Theory

The concepts and theories employed in this thesis are based in econometrics and macroeconomics. Consequently, the theoretical chapter will largely consist of the different important econometric concepts necessary to grasp this thesis. In section 3.1, the type of variables exhibited in the data; time series variables, will be explained. Nonstationary variables will also be introduced. Section 3.2 will comprise the concept of unit root- and stationarity tests, while cointegration of nonstationary variables will be covered in section 3.3.

## 3.1 Time Series Variables

In the field of econometrics, data collected on a single unit over time is referred to as a *time series* variable. Time series variables are dynamic in nature in that their current values will be correlated with their past values. Equivalently, their current values can also be related to current and past values of other time series variables. To account for this dynamic relationship, lagged values of certain variables can be included when modelling (Hill *et al.*, 2018). Examples of lagged values of different variables, where $t$ indexes time, and *q,p,s* indexes the number of lagged periods, are presented in Table 3.

TABLE 3: Lagged Values of Different Variables

| | Type of variable |
|---|---|
| Explanatory variable | $(x_{t-1}, x_{t-2}, ..., x_{t-q})$ |
| Dependent variable | $(y_{t-1}, y_{t-2}, ..., y_{t-p})$ |
| Error term | $(e_{t-1}, e_{t-2}, ..., e_{t-s})$ |

An important aspect of time series variables is the concept of *stationarity*. A stochastic process is stationary if its probability distribution is invariant to time (Pesaran, 2015). This implies that its mean and variance will be constant and time-variant. Hence, doing estimation on subsets of observations corresponding to different windows of time, we would estimate equal population quantities, mean $\mu$, variance $\sigma^2$ and autocorrelations $\rho_1, \rho_2, \rho_3, ...,$. These requirements are listed in equations 3.1a-c.

$$E(y_t) = \mu \qquad \text{(constant mean)} \qquad (3.1a)$$

$$\text{var}(y_t) = \sigma^2 \qquad \text{(constant variance)} \qquad (3.1b)$$

$$\text{cov}(y_t, y_{t+s}) = \text{cov}(y_t, y_{t-s}) = \gamma_s \qquad \text{(covariance depends on } s \text{ not } t\text{)} \qquad (3.1c)$$

When these requirements are violated, it is likely that the variables exhibit *nonstationary* features. These variables can have a non-constant mean, or the variance could be increasing with time, for example. Proceeding with a regression of these variables could potentially introduce the problem of *spurious regression*, where the regression results could falsely report strong linear relationships and significant coefficients.

## 3.2 Unit Root Tests

In order to assess the stationarity or nonstationarity of a variable, one should be familiar with the concept of *unit roots*. To appreciate the relationship between nonstationarity and unit roots, we begin with the autoregressive model presented in Dickey & Fuller (1979).

$$y_t = \rho y_{t-1} + e_t, \qquad t = 1, 2, ..., \tag{3.2}$$

$\rho$ is a real number and $e_t$ is the error term[4] with mean zero and variance $\sigma^2$. When $|\rho| < 1$, the autoregressive series converges to a stationary time series. When $|\rho| = 1$, the series is referred to as a *random walk* model and will slowly wander up and down with no discernible pattern. Means of subsamples of observations, in this case, would be dependent on the sample period, which is a nonstationary characteristic. In the case of $|\rho| > 1$, the series would be nonstationary and the variance would grow exponentially with time; the series is nonstationary and *explosive*[5]. When $|\rho| = 1$, $y_t$ has a unit root and the series is nonstationary[6]. To test the stationarity of a series, one can employ stationarity- and unit root tests. In this thesis, the *Augmented Dickey-Fuller* test (Dickey & Fuller, 1979) and the *Kwiatkowski-Phillips-Schmidt-Shin* test (Kwiatkowski *et al.*, 1992) will be employed.

### 3.2.1 Dickey-Fuller Test

The Dickey-Fuller test (DF test) is a hypothesis test where the null hypothesis is the existence of a unit root in a series $y$. The DF test can be specified according to different formulations of the alternative hypothesis. Specifically, there are three versions of the DF test: the alternative is stationary around a nonzero mean; the alternative is stationary around a linear trend; the alternative is stationary around a zero mean. Hence the DF test can assess different types of models:

[4] In terms of time series, the error terms are often referred to as "shocks" (Hill *et al.*, 2018).

[5] Due to its empirically rare occurrence, the case of $|\rho| > 1$ will not be further considered in this thesis.

[6] See section 12.3.1 in Hill et al. (2018) for further detailing on the origin of the unit root.

$$y_t = \alpha + \lambda t + \rho y_{t-1} + e_t \tag{3.3a}$$

$$y_t = \alpha + y_{t-1} + e_t \tag{3.3b}$$

$$y_t = y_{t-1} + e_t \tag{3.3c}$$

Here, 3.3a is a model with an intercept and a linear trend, 3.3b is a random walk with drift model, and 3.3c is a simple random walk model. Dickey and Fuller (1979) also proposed the *augmented* Dickey-Fuller (ADF) test, which could handle more complicated sets of time series models by including $\Delta y_{t-p}$ in the test equation. There have been developed other unit root tests that account for some of the weaknesses imposed by the DF test: the Phillips-Perron test (PP test) suggests a non-parametric test statistic for the null hypothesis of a unit root that explicitly allows for weak dependence and heterogeneity of the error process (Phillips & Perron, 1988); the Elliot-Rothenberg-Stock test (ERS test) removes the constant/trend effects from the data and takes into account the serial correlation of the error term (Graham *et al.*, 1996); the Kwiatkowski-Phillips-Schmidt-Shin test (KPSS test) specifies stationarity in the null hypothesis (Kwiatkowski *et al.*, 1992). The KPSS test will be further detailed in the next section.

### 3.2.2 The KPSS Test

The KPSS test is a test of level and trend stationarity, where the null hypothesis is stationarity of the data. The formulation of the hypotheses is thus opposite compared to the ADF test, where the null hypothesis is the presence of a unit root in the data. In a response to the trend of unit root tests at the time failing to reject the null hypothesis of unit roots in many economic time series, Kwiatkowski *et al.* (1992) suggested switching the hypotheses to stationarity in the null, which had been rarely attempted in the literature. They consider the model

$$y_t = \xi t + r_t + \epsilon_t \tag{3.4}$$

where $r_t$ is a random walk specified as

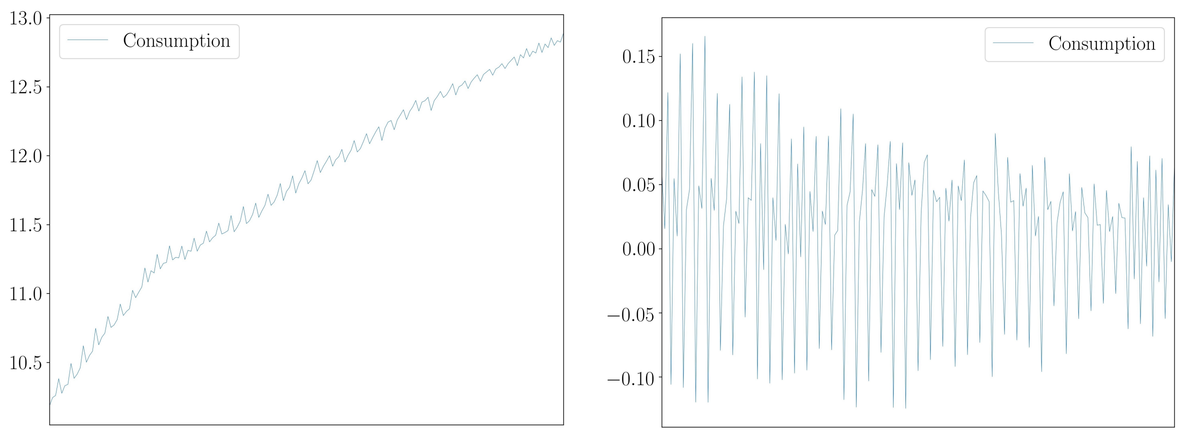
$$r_t = r_{t-1} + u_t \tag{3.5}$$

where $u_t$ is independent and identically distributed $(0, \sigma^2)$. The value $r_0$ serves as the intercept and corresponds to the level. Consequently, if $\xi = 0$, the model becomes $y_t = r_t + \epsilon_t$, and, under the null hypothesis $\epsilon_t$ is stationary, so $y_t$ would be level stationary. Conversely, if $\xi > 0$, $y_t$ is trend stationary. Additionally, the KPSS test is suited for conservative hypothesis testing, as the alternative is often the outcome of interest, and rejecting the null thus gives strong evidence of unit root presence (Pfaff, 2008).

### 3.2.3 Order of Integration

Some time series that exhibit nonstationary characteristics, can be made stationary by taking the first difference. A series $y_t$ is considered *difference stationary* if applying a first difference with respect to time results in a stationary process (Pfaff, 2008).

FIGURE 6: Normal and Differenced Data on Log of Final Consumption Expenditure of Households in Norway From 1978 to 2019




(a) (b)

*Source:* The data has been collected from SSB—StatBank Norway.

As an example, the log of the final consumption expenditure of households in Norway from 1978 to 2019 is plotted in Figure 6(a). Taking the first difference produces the plot in Figure 6(b), which from visual inspection seems to exhibit stationary characteristics. When taking the logarithm of such a series, one can eliminate the problem of increasing variance with time. Hence, when working with nonstationary variables, one should be aware of this possibility.

If a series becomes stationary after taking the first difference, it is considered to be integrated of order one, abbreviated as $I(1)$. More generally, the order of integration is often abbreviated as $I(d)$, where $d$ refers to the number of times the series must be differenced until it is stationary (Pfaff, 2008). If we take the first difference of equation 3.2, the result is

$$\Delta y_t = y_t - y_{t-1} = e_t \tag{3.6}$$

and considering the stationarity of the independent $(0, \sigma_v^2)$ random variable $e_t$, the first difference operation on $y_t$ makes it integrated of order one, or $I(1)$.

## 3.3 Cointegration

The concept of *cointegration* was first introduced by Granger (1981) and was properly formalized in Engle and Granger (1987). Two nonstationary $I(d)$ variables are said to be cointegrated if there exists a linear combination between them that yields a variable with a lower order of integration (Pfaff, 2008). In other words, if $y_t$ and $x_t$ are nonstationary $I(1)$ variables, and a linear combination of them, such as $e_t = y_t - \beta_1 - \beta_2 x_t$, turns out to be a stationary $I(0)$ process, $y_t$ and $x_t$ are cointegrated (Hill *et al.*, 2018). In regards to economics, this concept would allow economists to detect stable long-run relationships between nonstationary variables. Examples can also be found in finance, where interest rates of different maturities are cointegrated, and in macroeconomics, where we can find cointegration in the purchasing power parity hypothesis (Pesaran, 2015). Further, errors, or deviations, from the long-run equilibrium is possible, but would be characterized by mean reversion back to its equilibrium.

## 3.4 Testing for Cointegration

When testing for cointegration one can specify a test for the bivariate case, and for a system with more than two equations. For a single equation approach, Engle and Granger (1981) suggest a two-step residual-based approach. For a multiple equations approach, Johansen (1991) considers a vector autoregressive (VAR) model in error-correction form and employs matrix operations to calculate the rank of the cointegrating vector to specify the number of cointegrating relationships. It should be noted that while the multiple equations approach allows for more than one cointegrating relationship, the single equation approach only allows for one cointegrating relationship. The following two sections will consider these approaches in more detail.

### 3.4.1 Testing for Cointegration: Single Equation Approach

The *two-step* approach for cointegration in single equation was proposed by Engle and Granger (1987). In essence, the test for cointegration in single equation, say, for $y_t$ and $x_t$, is a test on the stationarity of the residuals, $e_t = y_t - \beta_1 - \beta_2 x_t$. If the residuals are stationary, $y_t$ and $x_t$ are cointegrated (Hill *et al.*, 2018). The stationarity of the residuals can be tested using the Dickey-Fuller test for unit roots. Assuming stationarity in the residuals, one subsequently specifies an *error-correction* model (ECM). More formally, if we start off with the one lag autoregressive distributed lag (ARDL) model

$$y_t = \delta + \theta_1 y_{t-1} + \phi_0 x_t + \phi_1 x_{t-1} + v_t \tag{3.7}$$

where $v_t$ represent the errors. We assume that $y$ and $x$ are cointegrated. If we set $y_t = y_{t-1} = y$, $x_t = x_{t-1} = x$ and $v_t = 0$, we can obtain

$$y(1 - \theta_1) = \delta + (\phi_0 + \phi_1)x \tag{3.8}$$

After doing some algebra and applying the definitions $\beta_1 = \delta/(1 - \theta_1)$ and $\beta_2 = (\phi_0 + \phi_1)/(1 - \theta_0)$, we obtain the error-correction equation

$$\Delta y_t = -\alpha(y_{t-1} - \beta_1 - \beta_2 x_{t-1}) + \phi_0 \Delta x_t + v_t \tag{3.9}$$

This embeds the cointegrating relationship between $y$ and $x$ in a general ARDL framework (Hill *et al.*, 2018). The term $(y_{t-1} - \beta_1 - \beta_2 x_{t-1})$ shows the deviation or error from the long-run equilibrium, while $\alpha$ shows the correction to the deviation. Hence, following the dynamics of the model, if the error in the previous period is positive $y_{t-1} > (\beta_1 + \beta_2 x_{t-1})$, $y_t$ will fall and the change will be negative. Further, if the opposite is true, $y_{t-1} < (\beta_1 + \beta_2 x_{t-1})$, $y_t$ will increase and the change will be positive. It is important to note that the error-correction part $\alpha$ must always be negative to prevent deviations from the long-run equilibrium. Also, as noted by Pfaff (2008), in the case of two integrated $I(1)$ variables, *Granger causality* must exist in at least one direction.

### 3.4.2 Testing for Cointegration: Multiple Equation Approach

In the multiple equation approach proposed by Johansen (1991), the presence of more than one cointegrating relationship is allowed. Begin with a general *vector* autoregressive (VAR) model or order $p$.

$$\mathbf{y}_t = \sum_{i=1}^{p} \mathbf{A}_i \mathbf{y}_{t-i} + \mathbf{u}_t \tag{3.10}$$

For $p > 1$ this VAR model can be re-written in VEC form

$$\Delta \mathbf{y}_t = \prod \mathbf{y}_{t-1} + \sum_{i=1}^{p-1} \mathbf{\Gamma} \Delta \mathbf{y}_{t-i} + \boldsymbol{\epsilon}_t \tag{3.11}$$

where $\mathbf{y}_t$ is the vector of $k$ nonstationary series, and $\Delta \mathbf{y}_t$ is its first-differenced values. $\prod = -(\mathbf{I} - \mathbf{A}_1 - ... - \mathbf{A}_i)$ for $i = 1, ..., p-1$ is the coefficient matrix of the cointegrating relationships, and $\mathbf{\Gamma}$ is the coefficient matrix of the lags of the differenced values of $\mathbf{y}_t$. The matrix $\prod$ contains the cointegrating vector $\boldsymbol{\beta}$, and the *loading* matrix $\boldsymbol{\alpha}$, which determines the speed of adjustment to the long-run equilibrium.

To appreciate the role of the $\prod$ matrix, consider the fact that the individual components of $\mathbf{y}_t$ are at most $I(1)$ variables so that the left-hand side of the VEC model (3.11) is stationary. In order to maintain balance, the term $\prod \mathbf{y}_{t-1}$ must also be stationary. Hence, the values of the elements in $\prod$ must be such that the right-hand side becomes stationary (Pfaff, 2008).

If $\prod = \mathbf{0}$, there is no cointegrating relationship between the variables and the model reduces to a VAR model in first difference. This corresponds to the case (i) where the *rank* of $\prod$ (henceforth denoted as $r$) is equal to zero. In the case (ii) $r = k$, the deviations of $\mathbf{y}_t$ around the deterministic components are stationary, hence the data are $I(0)$ in levels and it would be fitting to estimate a VAR in levels. In the case (iii) $r < k$, the $k \times r$ matrices $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ with rank $r$ exist such that $\prod = \boldsymbol{\alpha}\boldsymbol{\beta}'$. Consequently, $\boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{y}_{t-p}$ is stationary and each column in $\boldsymbol{\beta}$ are the cointegrating vectors, or the long-run relationships between the series in $\mathbf{y}_t$ (Pfaff, 2008). Rank hypothesis tests are run sequentially in the following manner: $r = 0$, $r \leq 1$, $r \leq 2$, and so on. If the test fails to reject $r = 0$ but rejects $r \leq 1$, it concludes that there exists no more than one long-run relationship between the variables. The same logic follows for $r \leq 2$, and so on. The test statistic can be specified to report the maximum eigenvalue or the trace statistic. One can note that for the question of which test statistic to emphasize when they report differing results, Lütkepohl *et al.* (2001) concludes that the power of the trace test and maximum eigenvalue test are largely similar, but notes that the trace test has superior power in some situation when the sample size is low.[7]

# 4 Methodology

Section 4.1 will give a presentation and discussion of the provided dataset. Also, the relevant wage data variables will be subject to necessary transformations, whereas in section 4.2 tests for stationarity will be further described Section 4.3 will describe and specify the Johansen test to test for cointegration between the variables to identify any long-run relationships. Finally, provided a cointegrating relationship(s) exists, a vector error-correction model will be fitted.

## 4.1 Data

The necessary data for this thesis has been provided by Statistics Norway (SSB). The wage variables are presented as hourly wage rates and are reported as time series variables in the quarterly range spanning 1978Q1 to 2018Q4, which gives a total of 160 observations. The immigration rate in proportion to the working-age population is included, as well as quarterly CPI inflation rate. Table 4 gives definitions and notations of the included variables, where $t$ is the time index, and $I, P_u, P_r, C, W$ denotes the different sectors. For the wage variables, it should be noted that their lower-case letter counterparts denote a logarithmic transformation.

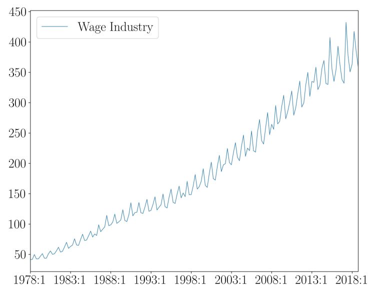
[7]See more in Johansen and Juselius (1990).

TABLE 4: List of Variables

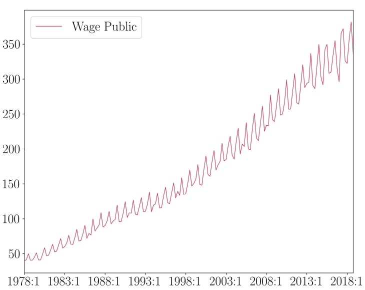
| Variables | Description |
|---|---|
| $HW_{It}$ | Hourly Wage Rate Industry Sector |
| $HW_{P_ut}$ | Hourly Wage Rate Public Sector |
| $HW_{P_rt}$ | Hourly Wage Rate Private Sector |
| $HW_{Ct}$ | Hourly Wage Rate Construction Sector |
| $HW_{Wt}$ | Hourly Wage Rate Wholesale Sector |
| $IM_t$ | Immigration Rate |
| $CPI_t$ | Quarterly CPI |

Figure 7a-e presents plots in levels of the hourly wage rates in the sectors relevant to this thesis. As is immediately apparent from visually inspecting the plots, they share many of the same properties.

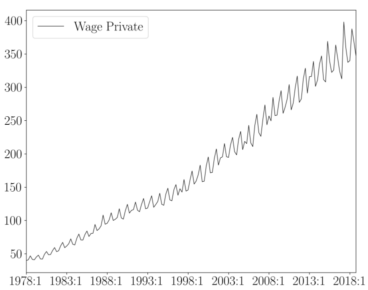
FIGURE 7: Plots of Hourly Wage Rates in the Industry, Public, Private, Construction and Wholesale Sectors



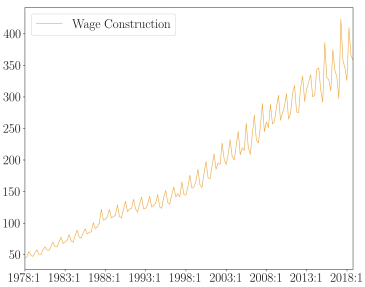

(a)



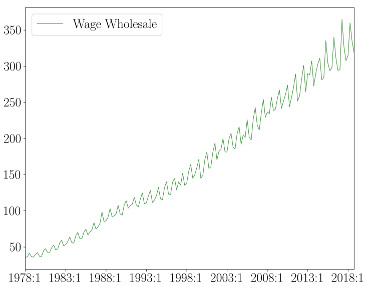

(b)




(c)




(d)



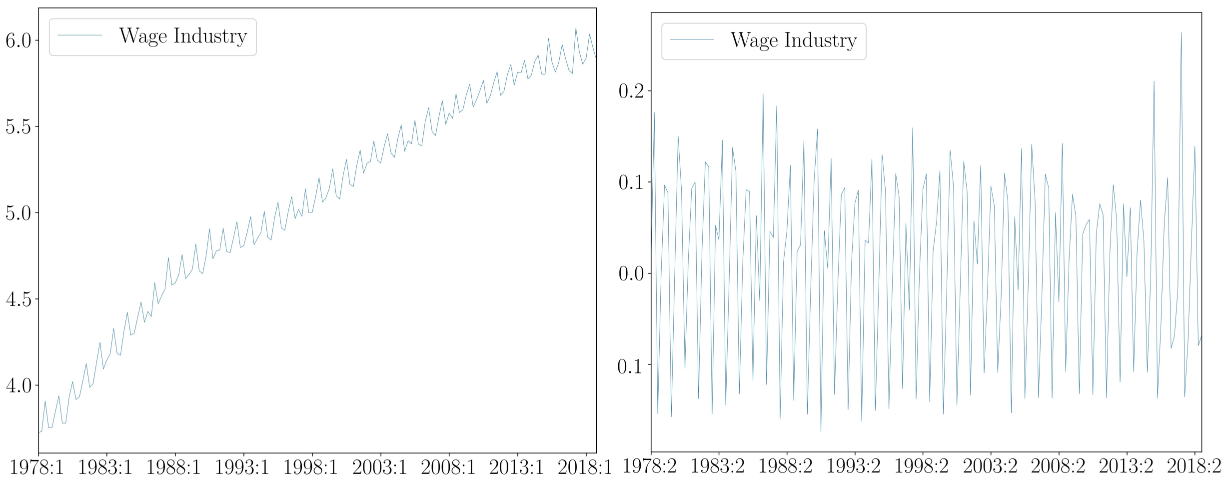

(e)

*Source:* The data has been collected from the KVARTS model variable list, provided by SSB.

In all sectors, there has been a marked increase in the hourly wage rates. Additionally, from visual inspection, all graphs do seem to exhibit nonstationary attributes, but this will require testing; see section 5.1. The variance in all wage variables appears to be increasing with time, this is accounted for by logarithmically transforming the variables.

Also, the increasing trend in the wage variables is resolved by taking the first difference. These transformations are sequentially performed and exemplified in Figure 8. As the wage variables seem to share similar statistical properties, it is sufficient to present these transformations on the industry sector as a visual representation; similar plots of the rest of the sectors are contained in the Appendix, Figure 10. In Figure 8a, a logarithmic transformation has been performed. The log-transformed variable is first differenced in Figure 8b, upon which seemingly visually stationary attributes are observed. This could potentially mean that the wage variables are $I(1)$.

FIGURE 8: Logarithmic and First Difference Transformations on the Hourly Wage Rate in Industry Sector

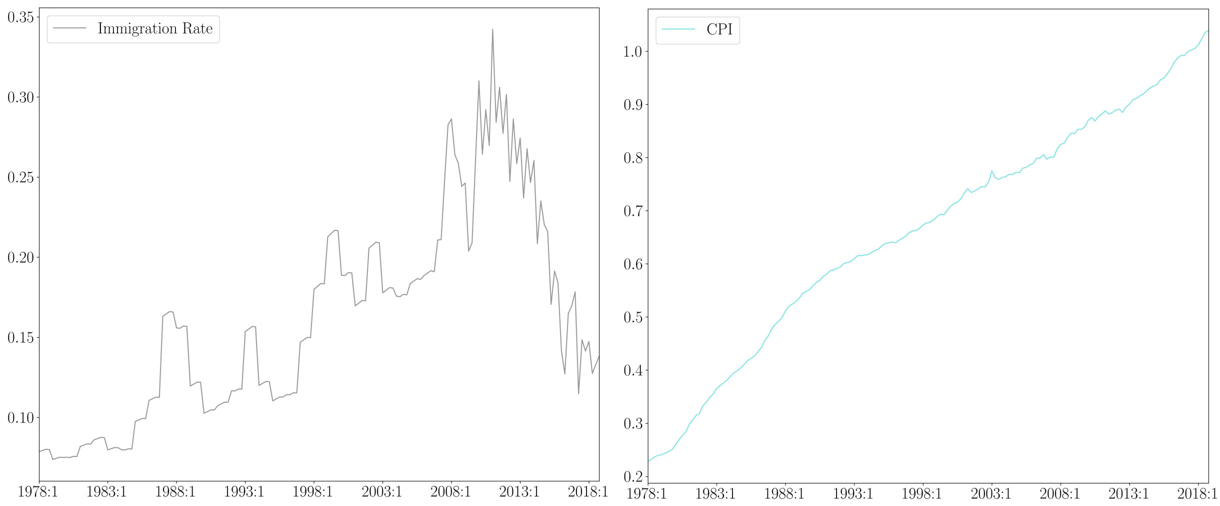



(a) (b)

*Source:* The data has been collected from the KVARTS model variable list, provided by SSB.

The immigration rate and CPI are visually presented in Figure 9. Until around 2012, the immigration rate seemed to be generally increasing, with some drops, for example after the 2008 financial crisis. After 2012, there has been a marked negative trend in the immigration rate throughout the remaining data sample period. In a report from Norway to the OECD, it is noted that the number of labour immigrants from non-Nordic countries has been on a decline since 2011 (Thorud, 2019).

FIGURE 9: Plots of the Immigration Rate and the CPI




*Source:* The data has been collected from SSB—StatBank Norway.

## 4.2 Stationarity Tests

In order to assess the stationarity of the wage data, both the ADF test and the KPSS test will be employed. As Pfaff (2008) notes in his book, there is rarely an altogether obvious choice of which tests should be applied to a data set. Hence, the pragmatic approach would be to include more than one test, preferably with opposing null hypotheses. The KPSS tests have been specified to test for trend-stationarity using a lag level of $\sqrt[4]{12 \times (n/100)}$. The lag length in the ADF test is 8, achieved through calculating the information criteria of the VAR process up to 10 lags. To further assess the order of integration of the wage variables, the ADF test will also be performed on the first differenced data.

## 4.3 Testing for Cointegration

The Johansen test will be employed to test for cointegration in the wage data. In the following, specification of the test will be detailed. The lag length included in the test is determined by assessing the information criteria as well as the final prediction error returned from sequentially increasing the lag order up to VAR($p$), $p = 10$ (Pfaff, 2005). A VAR is estimated by OLS per equation of the form

$$\mathbf{y}_t = A_1\mathbf{y}_{t-1} + ... + A_p\mathbf{y}_{t-p} + D_t + \mathbf{u}_t \tag{4.1}$$

where $\mathbf{y}_t$ is a $5 \times 1$ vector of the wage variables, $\mathbf{u}_t$ is an error term of equal dimension,

$A_p$ are $K \times K$ coefficient matrices and $D_t$ is a deterministic trend variable. The information criteria include the Akaike (AIC), the Hannan-Quinn (HQ), and the Schwarz (SC) information criteria, while FPE denotes the final prediction error. Their computations are listed below and the results are presented in Table 6.[8]

$$AIC(n) = \ln \det(\tilde{\sum}_u(n)) + \frac{2}{T} nK^2$$

$$HQ(n) = \ln \det(\tilde{\sum}_u(n)) + \frac{2\ln(\ln(T))}{T} nK^2$$

$$SC(n) = \ln \det(\tilde{\sum}_u(n)) + \frac{\ln(T)}{T} nK^2$$

$$FPS(n) = \left(\frac{T + n*}{T - n*}\right)^K \det(\tilde{\sum}_u(n))$$

TABLE 5: Results of Information Criteria Tests

| Lags | AIC(n) | HQ(n) | SC(n) | FPE(n) |
|---|---|---|---|---|
| $p_1$ | -37.80 | -37.56 | -37.20 | 3.84e-17 |
| $p_2$ | -39.38 | -38.94 | -38.29 | 7.89e-18 |
| $p_3$ | -40.53 | -39.88 | -28.94 | 2.51e-18 |
| $p_4$ | -42.62 | -41.77 | -40.53 | 3.13e-19 |
| $p_5$ | -42.52 | -41.47 | -39.93 | 3.49e-19 |
| $p_6$ | -42.62 | -41.36 | -39.53 | 3.20e-19 |
| $p_7$ | -42.48 | -41.03 | -38-90 | 3.72e-19 |
| $p_8$ | -42.61 | -40.95 | -38.53 | 3.35e-19 |
| $p_9$ | -42.52 | -40.66 | -37.95 | 3.76e-19 |
| $p_{10}$ | -42.39 | -40.33 | -37.31 | 4.47e-19 |
| $p_{11}$ | -42.23 | -39.97 | -36.67 | 5.46e-19 |
| $p_{12}$ | -42.40 | -39.93 | -36.33 | 4.91e-19 |

The reported lag lengths are calculated to be $AIC(n) = 6$, $HQ(n) = 4$, $SC(n) = 4$ and $FPE(n) = 4$. The selected lag length to be included in the Johansen test is determined to be 4.[9]

The immigration rate and CPI inflation are added to the test as exogenous regressors. The type of test is specified to report the *trace* statistic and a deterministic trend is

[8] $n*$ denotes the total number of parameters in each equation and $n$ denotes the lag order.

[9] It should be noted that the Johansen test does exhibit relatively high sensitivity to the selected lag length. This is further studied in i.e. Emerson (2007).

included in the cointegration relation. Additionally, the critical values are provided by Osterwald-Lenum (1992). It should be noted that, in regards to the thesis question posed in this paper: if the wage levels in the construction and wholesale sectors have followed the wage levels in the industry, public and private sectors, there will not be performed shock impulse responses on the variables. See Gjelsvik *et al.* (2015) for a relevant application of this analysis.

After running the Johansen test, a vector error-correction model can be fitted to extract the cointegrating vector and the short-run coefficients. The base vector autoregressive model the VEC model will be based on is expressed in the form presented in equation 6.1.

$$\Delta \mathbf{y}_t = \sum_{p=1}^{4} \mathbf{\Gamma}_p \Delta \mathbf{y}_{t-p} + \prod \mathbf{y}_t + \mathbf{x}\Theta_t + \epsilon_t \tag{4.2}$$

Here, $\mathbf{y}_t$ contains the vector $(\mathrm{HW}_i, \mathrm{HW}_{p_u}, \mathrm{HW}_{p_r}, \mathrm{HW}_c, \mathrm{HW}_w)$, while the exogenous variables IM and CPI are included in $\Theta_t$. Additionally, the parameters for the short-run and long-run effects are captured in the matrices $\mathbf{\Gamma}$ and $\prod$, respectively.

# 5 Results and Discussion

In section 5.1, the results of the unit root- and stationarity tests, explained in the methodology section, will be presented. Further, section 5.2 will disclose the cointegration test results, while the accompanying vector error-correction model is estimated in section 5.3. The model estimates will disclose the long-run and short-run effects present in the cointegrating relationship.

## 5.1 Unit Root- and Stationarity Tests

The results of the KPSS test rejects the null hypothesis of trend-stationarity in all wage variables. The results, which are presented in Table 6, thus indicate with high probability the presence of a unit root in the data.

TABLE 6: Results of the KPSS-Test

| Test Statistic | | | | | Critical Values | | |
|---|---|---|---|---|---|---|---|
| $\hat{\eta}_{\tau i}$ | $\hat{\eta}_{\tau p_u}$ | $\hat{\eta}_{\tau p_r}$ | $\hat{\eta}_{\tau c}$ | $\hat{\eta}_{\tau w}$ | 1% | 5% | 10% |
| 0.248 | 0.214 | 0.280 | 0.238 | 0.259 | 0.176 | 0.146 | 0.119 |

To further evaluate the nonstationarity of the data, the ADF test is performed. The test statistics for all variables, except for the public sector, fail to reject the null hypothesis, see Table 7. Regardless of the rejection of the null for the public sector, the results of the KPSS test, and the failure to reject the null for the remaining wage variables, imply the inclusion of the variable in subsequent tests. The ADF test is also performed on the first differenced data, which strongly rejects the null hypothesis of nonstationarity, see Table 8. Hence, it is statistically coherent to conclude that the variables are integrated of order one, $I(1)$.

TABLE 7: Results of the ADF-Test

| Test Statistic | | | | | Critical Values | | |
|---|---|---|---|---|---|---|---|
| $\tau_{3i}$ | $\tau_{3p_u}$ | $\tau_{3p_r}$ | $\tau_{3c}$ | $\tau_{3w}$ | 1% | 5% | 10% |
| -3.174 | -2.924 | -3.631 | -4.414 | -3.714 | -3.99 | -3.43 | -3.13 |

TABLE 8: Results of the ADF-Test on the First Differenced Wage Data

| Test Statistic | | | | | Critical Values | | |
|---|---|---|---|---|---|---|---|
| $\tau_{3\Delta i}$ | $\tau_{3\Delta p_u}$ | $\tau_{3\Delta p_r}$ | $\tau_{3\Delta c}$ | $\tau_{3\Delta w}$ | 1% | 5% | 10% |
| -19.201 | -18.978 | -21.171 | -19.184 | -20.592 | -3.99 | -3.43 | -3.13 |

The results of the stationarity tests support the nonstationarity of the wage variables. This is to be expected, as in the plots of the variables there is a visible increasing trend with time; a characteristic of nonstationary variables. Additionally, due to the nonstationarity of the variables, the Johansen test is now employed in order to find a linear combination(s) that results in a lower order of integration.

## 5.2 Cointegration Results

In the Johansen test for cointegration, the test statistic on the presence of a long-run relationship, $r = 0$, is greater than the critical values on the 10% and 5% level of significance. The test sequentially tests for more relationships, and the results imply that the trace test is unable to reject the null hypothesis of $r \leq 1$. Thus, the evidence does not support the presence of more than one cointegrating long-run relationship between the wage variables. In Table 9, the results of the Johansen test is presented. The $r$ denotes the *rank* of the cointegration vector.

TABLE 9: Results of the Johansen Test

| | Test Statistic | 1% | 5% | 10% |
|---|---|---|---|---|
| $r \leq 4$ | 5.59 | 16.26 | 12.25 | 10.49 |
| $r \leq 3$ | 12.39 | 30.45 | 25.32 | 22.76 |
| $r \leq 2$ | 28.34 | 48.45 | 42.44 | 39.06 |
| $r \leq 1$ | 54.90 | 70.05 | 62.99 | 59.14 |
| $r = 0$ | 94.69 | 96.58 | 87.31 | 83.20 |

The results are aligned with the dynamics of the main-course model; where Aukrust's wage-corridor maintains a stable long-run wage formation with the aid of the self-correcting mechanisms mentioned in section 2.1. As the cointegration analysis in Gjelsvik *et al.* (2015) provides a long-run relationship between the industry, public and private sectors, the evidence thus suggests that including wage levels in sectors that are suspected to have been particularly affected by the EU expansion, construction and wholesale, still constitute a cointegrating relationship. Additionally, despite not being formally tested after the inclusion of the construction and wholesale sectors, the results likely support the conclusion in Gjelsvik *et al.* (2015) of invariance of inflation targeting in regards to the dynamics of the main-course model. The results are also in line with other literature on the topic of cointegration in the framework of Norwegian wage formation, such as Nymoen (1989) and Eitrheim and Nymoen (1991). However, a potentially interesting matter of discussion is the marked drop in the immigration rate from 2011. In a 2017 SSB article, it is noted that in 2016, the net immigration was

26.100, the lowest net immigration seen since 2006. This sort of behaviour, increasing and decreasing in relatively equal magnitudes around a constant mean, could suggest that the immigration rate is stationary in nature.

## 5.3 Vector Error-Correction Model

The VEC model, estimated with four lags and one cointegrating vector, is contained in Table 12. while the loading matrix $\boldsymbol{\alpha}$ and the cointegrating vector $\boldsymbol{\beta}$ is presented in Table 10 and 11, respectively. The cointegrating vector contains the coefficients of which the long-run relationship is described; it is important to note that the cointegrating vector is normalized on the industry sector. The error-correction terms describe how quickly the wage variables converge back to the long-run equilibrium. These terms correspond to the coefficients of the $\boldsymbol{\alpha}$ vector; the loading matrix. So $\boldsymbol{\alpha}$, including the error-correction terms for the public, private, construction and wholesale sectors, respectively, becomes (-0.186,-0.023,-0.182,-0.286,-0.117). See Table 10. The results indicate that the error-correction terms of the industry and private sectors are significant on the 10% level, whereas the error-correction term for the construction sector is significant on the 5% level. The error-correction terms of the public and wholesale sectors are reported to be insignificant. The signs of the significant error-correction terms are correctly reported, and short-term deviations are corrected back to the long-run equilibrium. As mentioned, the error-correction term for the industry sector, the dependent variable, is significant on the 10% level and is reported to have the correct sign. The magnitude of the error-correction term reveals the time it takes to revert to the long-run equilibrium. The correction process is relatively quick, as short-run deviations are corrected back to the long-run equilibrium in around one year[10]. This would seem logical as the wage bargaining system in the Norwegian wage model incorporates yearly wage negotiations.

[10] $\frac{1}{4} * \frac{-1}{\ln(1-0.186)}$

TABLE 10: Loading Matrix $\boldsymbol{\alpha}$

| Variable | $HW_{it}.d$ | $HW_{p_ut}.d$ | $HW_{p_rt}.d$ | $HW_{ct}.d$ | $HW_{wt}.d$ |
|---|---|---|---|---|---|
| Estimate | -0.186 | -0.023 | -0.182 | -0.286 | -0.117 |
| Std.error | (0.104) | (0.107) | (0.095) | (0.112) | (0.097) |
| $t$-statistic | -1.792 | -0.219 | -1.917 | -2.555 | -1.209 |
| Pr(>\|t\|) | 0.075 | 0.827 | 0.057 | 0.012 | 0.229 |
| | · | | · | * | |

Significance codes: '·' 0.1, '*' 0.05, '**' 0.01, '***' 0.001

The cointegrating vector in Table 11 is the $\boldsymbol{\beta}$ vector, hence $\boldsymbol{\beta'}$ = (1.000,-4.590,7.312,-0.482,-3.545). We can look closer at the long-run equilibrium by solving for the dependent variable, wage in the industry sector, and removing the short-run effects. After solving for $HW_{it}$, the result is equation 5.1. The wage in the public, private, and wholesale sectors are all significant down to the 0.1% level, whereas the wage in the construction sector is barely significant on the 10% level. The trend variable, however, is insignificant. A one percent permanent increase in the hourly wage rate in the public sector would lead to a 4.590 percent increase in the hourly wage rate in the industry. An equal permanent increase in the hourly wage rate in the private sector, however, would lead to a 7.362 percent decrease in the hourly wage rate in the industry sector. One percent increases in the construction and wholesale sectors lead to an increase of 0.482 percent and 3.545 percent in the industry hourly wage rate, respectively.

$$\mathrm{HW}_{it} = 4.590\mathrm{HW}_{p_ut} - 7.362\mathrm{HW}_{p_rt} + 0.482\mathrm{HW}_{c_rt} + 3.545\mathrm{HW}_{w_rt} - 0.001t \tag{5.1}$$

TABLE 11: Cointegrating Vector $\boldsymbol{\beta}$

| | ect1 | Std.error | $t$-statistic | Pr(>\|t\|) | |
|---|---|---|---|---|---|
| $HW_{it}$ | 1.000 | | | | |
| $HW_{p_ut}$ | -4.590 | (0.786) | 5.841 | 8.3e-08 | *** |
| $HW_{p_rt}$ | 7.362 | (1.487) | 4.951 | 4.5e-06 | *** |
| $HW_{ct}$ | -0.482 | (0.285) | 0.012 | 0.096 | · |
| $HW_{wt}$ | -3.545 | (0.660) | -5.373 | 7.2e-07 | *** |
| trend | 0.001 | (0.001) | 0.573 | 0.338 | |

Significance codes: '·' 0.1, '*' 0.05, '**' 0.01, '***' 0.001

In Table 12, the short-run lagged differenced variables are listed. The lagged industry wage parameters are consistently highly significant in every lag, whereas we see significant parameters in the second and third lags of the construction-, and the first and second lags of the wholesale sectors. The short-run effect of the CPI on the wage in the industry sector is significant, which is in line with Bårdsen *et al.* (2005), where it is stated that the main-course theory does not rule out significant short-run effects of the CPI in a dynamic wage equation. Gjelsvik *et al.* (2015) estimates two simultaneous equations models, one before, and one after the implementation of an inflation targeting monetary policy, to try to capture the invariance of the CPI. Their results are that CPI expectations carried an effect on the wage system in both models. A result that could be expected as CPI expectation was always important in the Norwegian wage formation model, and is in line with the significant CPI effect in the VEC model. Following from Table 12, we see that the hourly wage rate in the industry sector is negatively affected by its first, second, and third lags, positively affected by the first and second lag of the wholesale sector, and negatively affected by the second and third lag of the construction sector. We also see that the public and sector sectors have no significant effect on the wage in the industry sector. In the Appendix, tabulated results of the VEC models where the the dependent variables are specified to be the public and private sectors are presented. When the public sector is the dependent variable, significant effects are noted in the first and third lags of the industry sector. Similarly, when the private sector is specified to be the dependent variable, we see significant effects on the first, second, and third lag of the wage in the industry sector. These results are in line with the fact that the wage in the industry sector exhibit a wage leader role in the Norwegian model. These results are also in line with Gjelsvik *et al.* (2015), where the industry, public, and private sectors were included. Due to limited litarature on the effects of wages in undersectors on the industry sector, it is difficult to compare and contrast the short-run effects of the construction and wholesale sectors with other results, hence interpretation would be mostly speculation.

TABLE 12: Estimated VEC Model with Cointegration Rank $r = 1$

| Variable | $HW_{it}.d$ | Std.error | $t$-statistic | Pr(>\|t\|) | |
|---|---|---|---|---|---|
| **Error-correction terms** | | | | | |
| ect1 | -0.186 | (0.104) | -1.792 | 0.075 | · |
| **Deterministic** | | | | | |
| Constant | -0.015 | (0.056) | -0.283 | 0.778 | |
| **Lagged differences** | | | | | |
| $IM_t$ | 0.158 | (0.072) | 2.200 | 0.029 | $*$ |
| $CPI_t$ | -0.155 | (0.045) | -3.458 | 7.2e-04 | $***$ |
| $HW_{it}.dl1$ | -1.045 | (0.306) | -3.420 | 8.2e-04 | $***$ |
| $HW_{p_ut}.dl1$ | 0.012 | (0.190) | 0.061 | 0.952 | |
| $HW_{p_rt}.dl1$ | -0.184 | (0.625) | -0.294 | 0.769 | |
| $HW_{ct}.dl1$ | -0.321 | (0.204) | -1.575 | 0.119 | |
| $HW_{wt}.dl1$ | 0.667 | (0.327) | 2.037 | 0.044 | $*$ |
| $HW_{it}.dl2$ | -1.088 | (0.289) | -3.766 | 2e-04 | $***$ |
| $HW_{p_ut}.dl2$ | 0.084 | (0.289) | 0.290 | 0.773 | |
| $HW_{p_rt}.dl2$ | -0.171 | (0.786) | -0.218 | 0.828 | |
| $HW_{ct}.dl2$ | -0.428 | (0.239) | -1.790 | 0.076 | · |
| $HW_{wt}.dl2$ | 0.789 | (0.391) | 2.018 | 0.045 | $*$ |
| $HW_{it}.dl3$ | -1.210 | (0.298) | -4.061 | 8.1e-05 | $***$ |
| $HW_{p_ut}.dl3$ | 0.206 | (0.382) | 0.539 | 0.591 | |
| $HW_{p_rt}.dl3$ | 0.823 | (0.882) | 0.934 | 0.352 | |
| $HW_{ct}.dl3$ | -0.495 | (0.209) | -2.372 | 0.019 | $*$ |
| $HW_{wt}.dl3$ | 0.163 | (0.457) | 0.358 | 0.721 | |

Significance codes: '·' 0.1, '*' 0.05, '**' 0.01, '***' 0.001

# 6 Conclusion

This thesis investigated whether the wage levels in the construction and wholesale sectors have followed those of the industry, public, and private sectors, since the expansion of the EU into Eastern Europe and the implementation of inflation targeting monetary policy. In order to ensure that the wage variables were applicable for a test for cointegration, the nonstationarity of the variables was assessed using unit root- and stationarity tests. A multivariate cointegration test results in the existence of one cointegrating vector, implying that there is a long-run equilibrium relationship between the wage variables. Moreover, the results also add to the conclusions in Gjelsvik *et al.* (2015) in that the Norwegian model continues to ensure that the wage levels in different sectors in the Norwegian economy grow towards a long-run equilibrium, with the sector-wise wage levels never straying too far off the cointegrating relationship. Furthermore, the short-run effects of the estimated VEC model adds to the wage leader dynamic of the wage in the industry sector.

There was an assessment regarding the inclusion of the immigration rate and the CPI as exogenous variables in the Johansen test, in favor of a more focused, and less general model, only focusing on the wage variables. Including the two exogenous variables in the test created a more general model, and potentially a better-specified test for cointegration between the wage variables. Furthermore, specifying a perfect econometric model is difficult, so the pragmatic approach becomes conveying justification and explanation for the inclusion of certain variables.

Looking more closely at the results in light of Gjelsvik *et al.* (2015), more evidence has been collected through testing to establish the invariance of the structural changes to the cointegrating dynamics of the Norwegian model, despite their findings that the wage level in the industry sector was reduced as a result of the negative and significant effect of the increased labour immigration. Thus, to answer the question posed in the introductory section, the implications seem to be that the main-course model, despite having endured several structural changes, the EU expansion and inflation targeting included, continues to be a sturdy and stable model for wage formation. However, questions can be raised regarding the improvement of unionizing labour immigrants, and if this would have worked to neglect the negative effect on the wage level in the industry sector.

It is worth noting that, in the future, there are ample opportunities to study the Norwegian model in light of recent developments in the labour market. In a 2018 NHO report, it is stated that digitalization and automation can impose lower demand for low-education jobs. In the same report, a projection of the demand for skills up to 2035 showed that the largest source of demand would arise from skills procured in bachelor's degrees and vocational education. Conversely, the number of highly educated individuals in Norway is increasing, potentially causing a trend of overqualified workers in lower-educated jobs. Hence, it is reasonable to believe that there is a probability that the Norwegian model will be challenged by other structural changes, taking into account a rapidly changing labour market.

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# Appendix

TABLE 13: Estimated VEC Model with Cointegration Rank $r = 1$. Dependent Variable: *Public Sector*.

| Variable | $HW_{put}.d$ | Std.error | $t$-statistic | Pr(>\|t\|) | |
|---|---|---|---|---|---|
| **Error-correction terms** | | | | | |
| ect1 | -0.023 | (0.107) | -0.219 | 0.827 | |
| **Deterministic** | | | | | |
| Constant | 0.062 | (0.056) | 1.100 | 0.273 | |
| **Lagged differences** | | | | | |
| $IM_t$ | 0.089 | (0.074) | 1.195 | 0.234 | |
| $CPI_t$ | -0.079 | (0.046) | -1.708 | 0.089 | · |
| $HW_{it}.dl1$ | -0.715 | (0.315) | -2.268 | 0.024 | $*$ |
| $HW_{put}.dl1$ | -1.014 | (0.197) | -5.156 | 8.4e-07 | $***$ |
| $HW_{prt}.dl1$ | 0.772 | (0.645) | 1.196 | 0.234 | |
| $HW_{ct}.dl1$ | -0.114 | (0.210) | -0.541 | 0.560 | |
| $HW_{wt}.dl1$ | 0.246 | (0.338) | 0.731 | 0.466 | |
| $HW_{it}.dl2$ | -0.366 | (0.298) | -1.228 | 0.222 | |
| $HW_{put}.dl2$ | -1.229 | (0.298) | -4.124 | 6.3e-05 | $***$ |
| $HW_{prt}.dl2$ | 1.208 | (0.811) | 1.490 | 0.139 | |
| $HW_{ct}.dl2$ | -0.532 | (0.247) | -2.155 | 0.033 | $*$ |
| $HW_{wt}.dl2$ | 0.126 | (0.404) | 0.311 | 0.756 | |
| $HW_{it}.dl3$ | -0.739 | (0.308) | -2.401 | 0.017 | $*$ |
| $HW_{put}.dl3$ | -1.023 | (0.395) | -2.592 | 0.010 | $*$ |
| $HW_{prt}.dl3$ | 2.258 | (0.911) | 2.480 | 0.014 | $*$ |
| $HW_{ct}.dl3$ | -0.548 | (0.216) | -2.538 | 0.012 | $*$ |
| $HW_{wt}.dl3$ | -0.493 | (0.471) | -1.046 | 0.297 | |

Significance codes: '·' 0.1, '*' 0.05, '**' 0.01, '***' 0.001

Notes: Setting the public sector as the dependent variable, we find that the first and the second lags of the wage in the industry sector induce significant short-run effects on the wage in the public sector.

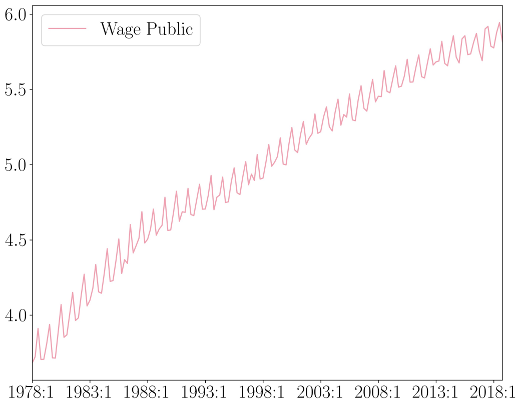
TABLE 14: Estimated VEC Model with Cointegration Rank $r = 1$.
Dependent Variable: *Private Sector*.

| Variable | $HW_{p_r t}.d$ | Std.error | $t$-statistic | Pr(>\|t\|) | |
|---|---|---|---|---|---|
| **Error-correction terms** | | | | | |
| ect1 | -0.182 | (0.095) | -1.917 | 0.057 | · |
| **Deterministic** | | | | | |
| Constant | -0.013 | (0.049) | -0.263 | 0.793 | |
| **Lagged differences** | | | | | |
| $IM_t$ | 0.142 | (0.066) | 2.151 | 0.033 | $*$ |
| $CPI_t$ | -0.152 | (0.041) | -3.697 | 3.1e-04 | $***$ |
| $HW_{it}.dl1$ | -0.465 | (0.279) | -1.664 | 0.098 | · |
| $HW_{p_u t}.dl1$ | 0.083 | (0.174) | 0.479 | 0.633 | |
| $HW_{p_r t}.dl1$ | -0.897 | (0.571) | -1.570 | 0.119 | |
| $HW_{ct}.dl1$ | -0.194 | (0.186) | -1.044 | 0.298 | |
| $HW_{wt}.dl1$ | 0.619 | (0.299) | 2.069 | 0.040 | $*$ |
| $HW_{it}.dl2$ | -0.455 | (0.264) | -1.722 | 0.087 | · |
| $HW_{p_u t}.dl2$ | 0.014 | (0.264) | 0.053 | 0.957 | |
| $HW_{p_r t}.dl2$ | -0.785 | (0.718) | -1.093 | 0.276 | |
| $HW_{ct}.dl2$ | -0.351 | (0.218) | -1.610 | 0.109 | |
| $HW_{wt}.dl2$ | 0.797 | (0.358) | 2.228 | 0.027 | $*$ |
| $HW_{it}.dl3$ | -0.789 | (0.272) | -2.897 | 0.004 | $**$ |
| $HW_{p_u t}.dl3$ | 0.308 | (0.349) | 0.882 | 0.379 | |
| $HW_{p_r t}.dl3$ | 0.086 | (0.806) | 0.107 | 0.915 | |
| $HW_{ct}.dl3$ | -0.379 | (0.191) | -1.984 | 0.049 | $*$ |
| $HW_{wt}.dl3$ | 0.287 | (0.418) | 0.686 | 0.493 | |

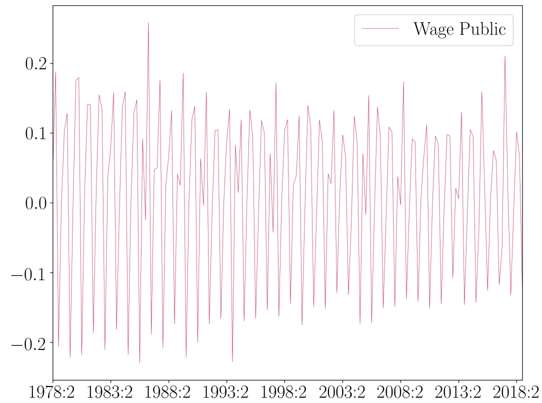
Significance codes: '·' 0.1, '*' 0.05, '**' 0.01, '***' 0.001

Notes: Setting the private sector as the dependent variable, we find that the first and the second lags of the wage in the industry sector induce significant short-run effects on the wage in the public sector.

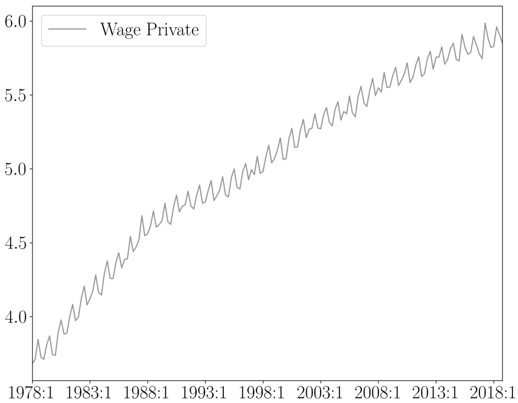
FIGURE 10: Logarithmic and First Difference Transformations on the Hourly Wage Rate in the Public, Private, Construction, and Wholesale Sectors



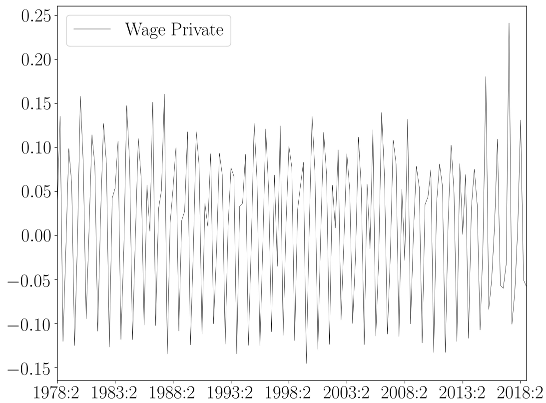

(a) Log Public



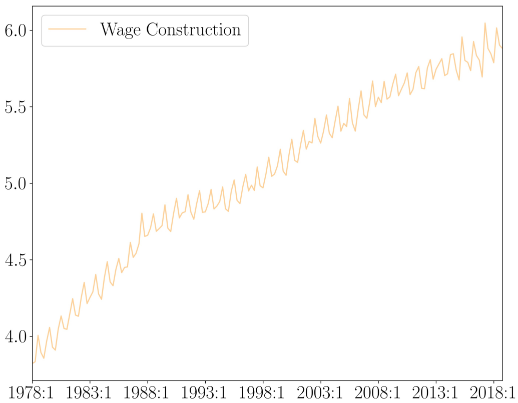

(b) Log Differenced Public



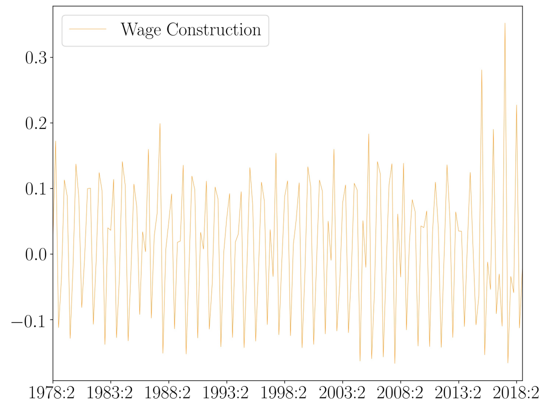

(c) Log Private



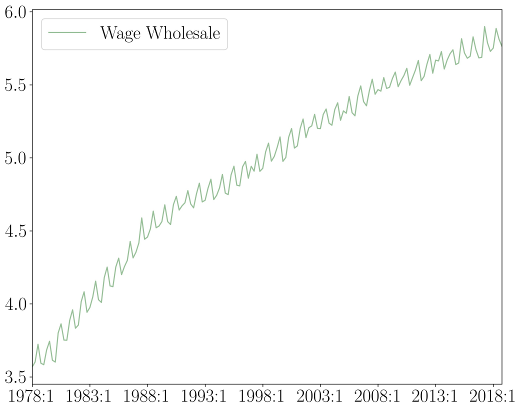

(d) Log Differenced Private



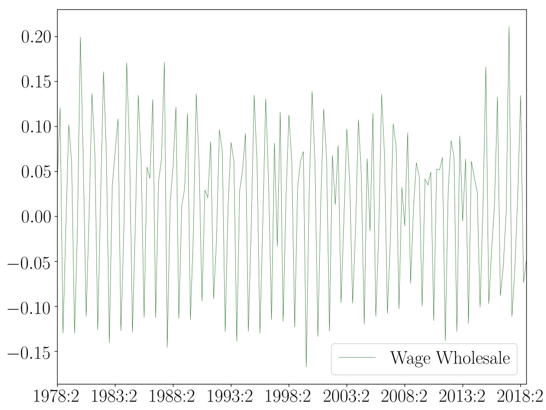

(e) Log Construction




(f) Log Differenced Construction




(g) Log Wholesale




(h) Log Differenced Wholesale

*Source:* The data has been collected from the KVARTS model variable list, provided by SSB.