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Variable and Norwegian Wage Formation, 1900–2025**

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From Manufacturing Productivity to National Wages: The Main-Course Variable and Norwegian Wage Formation, 1900–2025.*

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Abstract

We analyse the role of the wage scope defined by the manufacturing sector in the history of wage coordination in Norway. Using econometric methods, we document a long-term empirical equation that link the national annual wage to the manufacturing-sector wage scope, defined as the value of labour productivity, and which is known as the 'main-course'. Tests of exogeneity support the interpretation of the main-course as a long-run driver of the national annual wage. Contributing to the literature on the wage–productivity nexus, we show that the productivity component of the main-course has been the dominant determinant of the secular growth in the purchasing power of annual wages. Changes in the relative price of manufacturing value added and consumer goods have played a subsidiary role, together with unemployment and reductions in annual working time. Identified structural breaks cluster in the 1920s and 1930s, reflecting economic crises and institutional change, while the estimated relationship remains consistent with Norway's contemporary pattern wage bargaining system.

Keywords: Pattern wage bargaining; Wage-price nexus; Main-course model; Wage coordination; Long-run relationship; Cointegration; Economic history.

1 Introduction

Labour market institutions are comparatively strong in Norway and collective bargaining continues to be important in the wage formation processes. International comparisons place Norwegian together with countries that have “organised decentralised and coordinated bargaining systems”, OECD (2019a, p. 74).

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The current system in Norway operates through pattern wage bargaining, which assigns special importance to the manufacturing sector, Calmfors (2025). The peak organisations in this sector use profitability benchmarks to negotiate a wage growth target for the entire economy. A hallmark of pattern wage bargaining is the clear sequencing of decisions each year. The frontrunner is the agreement in the manufacturing industry.¹ After the frontrunner, sectoral agreements are negotiated in the wage following sectors of the economy.²

However, the sequencing of decisions in wage setting is not a new invention. It is more structural trait that links the modern pattern bargaining to earlier systems, in which the wage scope defined by the “tradable” (manufacturing) sector served as both an anchor and a driver of the general wage level. Another structural feature that connects the past with the wage formation of today is the institutions of resolution in the case of conflict about pay and working conditions, which can be traced back to the mid-1930s.

The institutionalisation of collective bargaining was vital for the labour movement during Norway’s industrialisation. It was a barrier against unregulated competition, sometimes referred to as “Manchester liberalism”, Frisch (1945). After a period marked by industrial unrest and class conflicts, it was also politically recognized as a tool for achieving relative peace between workers and employers. Labour market regulation had to be compatible with private ownership and competitive product markets. This balancing act, between liberalized product markets and labour markets protected just enough to sustain collective bargaining, shaped wage setting in Norway. It became a defining feature of labour market development in the Western economies during the 20th century.

Because economies and labour markets are evolving, the balancing act continues. From the perspective of the builder of empirical models of national wage formation, a key implication of the duality between institutional forces and market dynamics is that both have shaped the data over a long historical period. Consequently, despite high year-to-year variability, and secular changes in industry structure and labour market participation, there may be an element of continuity in the data-generating process, justifying econometric treatment to wage formation over a long time span.

This paper focuses on the degree of continuity, stability, and invariance in one particular institutional dimension of collective wage formation: the role and influence of the *main-course* variable, which originates in the manufacturing sector, in determining the trend of the national wage level.

The theoretical framework of wage setting with a clearly defined main-course variable dates back to the 1960s, see Bårdsen et al. (2005, Chapter 3) for a brief account. At that time, the manufacturing sector was about to reach its peak, in terms of employment share and as share of total value added in Norway. Given that leading role, it was perhaps only natural that collective agreements in manufacturing strongly influenced wage setting in other sectors. However, it is also plausible that the concept of the main-course was theoretical formalization of a practical wage-setting system with a longer history.

¹Head agreements (*hovedoppgjør*) on wages and broader issues last two years. The *frontrunner* is then the agreement between the Federation of Norwegian Industries (*Norsk Industri*) and two unions (*Fellesforbundet* and *Parat*). In the intermediate settlement (*mellomoppgjør*)—which covers only wage adjustments—the frontrunner is the agreement between peak organisations (NHO and LO, and NHO and YS).

²Sectoral agreements leave room for workplace-level agreements to finalize the wage contract, and the challenge is then to obtain sufficient coherence between the frontrunner wage target and the national wage increase generated by the two-tiered process. This trait is shared with other countries with decentralized coordination, Denmark, Finland, the Netherlands and Sweden, OECD (2019a, Chapter 2.7).

While it goes beyond this paper to discuss how to place the main-course theory's in the history of economics ideas, it is not too much of a speculation to claim that it was inspired by a certain exogeneity proposition about the wage level, Forder (2014, Ch. 1), that was found on in macroeconomic thinking and teaching in the first post World War II decades. It was not that the wage level was seen as unconditionally fixed or determined from outside. It was more the point that the wage level was recognized theoretically to be only loosely attached to the level of unemployment. Hence the meaning of exogeneity of nominal wages and prices was with respect to the level of unemployment (and other measures of supply-demand imbalance, like output-gap). In these models, it was seen as possible that wage and price growth could be stabilized at a given rate of (full) unemployment, which in turn may be seen a continuation of the discussion in (Keynes, 1936, Chapter 19).

Also mentioning is the conceptual closeness to the ability-to-pay hypothesis that gained some popularity in the 1950s and 1960s. As pointed out in Carruth and Oswald (1989, Chapter 1), research findings by leading economists in the 1950s and 1960s, had strongly indicated that compensation and pay levels depended on factors which were determinants of profits.³ Moreover, that form of division of value added between the worker and firms side was regularly and broadly taking place. As we review below, the same idea is found in the work of Scandinavian economists of the same era, The term main-course (of the wage level) was coined by Aukrust (1977). Conceptually the main-course corresponds to long-term ability-to-pay.

In this paper, we make use of new historical time series of prices, productivity and wages from the early 1900s to 2025, Harildstad (2026). This allows us to define a long time series for the main-course variable, and to empirically assess its force. We conceptualize force along three dimensions. First, we investigate the existence of a relationship between the main-course and the national wage, using graphical analysis and econometric methods. Second, we examine the stability of the equilibrium correction wage equation implied by a long-run relationship. Third, we address the empirical question of causality: Has the main-course acted as a driver of the national wage, or are there signs of reverse causality?

We contribute to the literature that give econometric treatment to aggregate wage formation in Norway by using cointegration methods and equilibrium correction models, see Nymoen (1989a, 1989b), Johansen (1997), Bårdsen et al. (2005) among others. Our finding that the productivity component of the main-course has been the main factor behind secular growth in the purchasing power of annual wages, contributes to the literature of wage-productivity nexus, widely defined, see Castle and Hendry (2009), Dosi et al. (2020), Schwellnus et al. (2024) among others. The constancy of the relationship, over the long-term, indicates that the Decoupling Phenomenon (the Productivity-Pay Gap) has not yet manifested itself macroeconomic data set that we analyse.

A paper thematically close to ours is Gjelsvik et al. (2020) who presented a model of pattern wage and price setting in Norway with quarterly data, but over a much shorter historical period than we span with our data set. Dalnoki (2020), estimated an econometric model with a wage-leading and a wage-following sector, using annual data from 1970 to 2018.

The rest of the paper is organised as follows: Section 2 gives a brief account of how the system of coordination has developed historically. We give attention to the adaption to changes in market conditions, industrial structure, economic policies, and

³Carruth and Oswald mention Slichter (1950) and Lester (1952) among others.

in the the relative power of labour market organisations. Section 3 introduces a simple theoretical framework with wage and price equations that imply logically a long-run relationship between the main-course variable and the national wage. Section 4 contains the econometric analysis of that hypothesized relationship. The final section summarizes the main findings and discusses their relevance for current debate about the robustness and functionality of Norway’s pattern wage bargaining system in our modern economy.

2 Lineages of wage coordination in Norway

In Norway, as elsewhere, labour market organisation is a phenomenon associated with industrialization, which happened late in Norway. The economy was little affected by the first industrial revolution. However, by 1900, demand for labour in shipbuilding, mining and factories powered by the country’s many waterfalls, had led to concentrations of wage earners in the bigger cities and in mining towns.

2.1 Organisation building

Industrial disputes (strikes and lockouts) about wage compensation, length of the working day and working conditions, had not been unknown during the second half of the 19th century. But in this era, the rule was that workers’ action had little basis in organisation. There were important initiatives and “movements”, but they were usually short-lived. For example, the *Thrane movement* was nation wide, and reached high member numbers in middle of the 1800s, but it was short-lived. Fragmentation, weak solidarity and downright rivalisation between workers’ organisations was also common, Bals (2022, Chapter 2 and 3).

Gradually, the ability to organise and to generate solidarity between worker’s became stronger. The Norwegian trade union confederation (LO) dates back to 1899. It was formed by trade unions that became central in the evolution of labour market reforms in Norway in the 20th century. Organisation of firms and capital owners developed as if in tandem. The peak organisation on the firm side was formed in 1900.⁴

Still, the organisation percentage for workers (union density) may have been below 10 percent in 1900, Olstad (2009). But it increased year by year, and reached 50 percent at the end of the 1930s. The number and coverage of collective agreements also increased both before World War I, and in the interwar years, Olstad (2009, p. 436-437).

Using data from the the 1930s in Norway, Bårdsen et al. (2010) presented evidence of a functional relationship between the manufacturing wage and factors that determine profitability and ability to pay.⁵ Theoretically interpreted, such a relationship is consistent with collective wage bargaining playing a major role in the wage setting process, Forslund et al. (2008).

2.2 Industrial strife and The Great Reconciliation

The growth of trade unions and the acceptance of collective bargaining did not happen without conflict. Strife was the normal situation. In particular, 1921, 1924 and 1931 were years with serious conflicts. According to historians, this was a dangerous period

⁴NAF, which is NHO nowadays.

⁵The data used by Bårdsen and Klovland was a panel data set of individual manufacturing firms.

for Norwegian society. Conflicts were so many, and fought so hard, that they became a real threat to political democracy.

1934 became a year of decision – about having a common destiny or not. It must have been surprising at the time that the parties chose reconciliation rather than mutual destruction, Olstad (2019). Then, in 1935, came the Master agreement between the two peak organisations LO and NAF. It has been regarded as the foundation of modern labour market regulation in Norway. Leiserson (1959) is an early example of an “onlooker’s” impression of the Master agreement in 1935 as a turning point: away from strife and towards a joint capacity for coordinated, concerted adjustments in several key areas.⁶

The development of public sector wage setting institutions followed its own path. Right until 1940, public sector wage setting was highly decentralised, Falch (2001). For example, teachers had no formal collective bargaining rights before the first agreement between the national teacher union and the (central) government was established in 1948, Falch and Strøm (2025) Nevertheless, the teacher wage relative to the manufacturing wage was almost the same in 1939 as in 1905, a reminder that medium term relative wage stability in itself is not evidence of (or depends on) formal centralization of wage formation.

2.3 Consolidation of collective wage setting and development of main-course theory

As the postwar period unfolded, with *de facto* full employment, and with a government committed to support free collective bargaining, the management of the economy in Norway became focused on the trade balance, exchange rate policies and international wage cost competitiveness. In Norway, as in many other countries, inflation became a problem for reaching important goals. It was not seen as an instrument towards attainment of those goals.⁷

Inflation was considered an integral part of the processes that governed the international cost competitiveness of Norway’s export and import competing industries. The role of wage formation in these processes was clearest conceptualized in a government commissioned report written by a group of experts in 1967. It soon became known as the Main-course model.⁸

A premise of the model was that the value added created in productive processes represented a rent that would be divided between capital owners and workers. Another was that the distribution of value added per hour occurs across both the internationally competitive sectors (henceforth, the tradables sector) and the sectors sheltered from international competition (the non-tradables sector). However, the strategic position of

⁶Reiersen (2024) is an interesting analysis of the mental re-orientation during the 1920s and 1930, which probably was needed on both sides in order to break the deadlock marked by strife and industrial unrest. Reiersen’s view is that the main step was to move from a situation of distrust, and hence conflict as the main strategy on both sides, to a situation with sufficient trust so that the mutual strategy became one of compromise and cooperation.

⁷Contrary to the ‘Phillips curve myth’ that formed between 1975 and 1977, there are almost no evidence of Phillips curve inflationism in Britain, as Forder (2014) shows convincingly. One can say the same about Norway, where inflation was unpopular among union leaders and members, Bergh (2009, p 118).

⁸Re-labelled as ‘A Norwegian Model of Inflation’ by Aukrust (1977), perhaps in an attempt by Aukrust, who had headed the 1967-group, to coin something different than the term ‘Scandinavian model’ after Edgren et al. (1969).

firms differs significantly between the two. In the non-tradables sector, firm owners can protect their share of value added by adjusting product prices after wage levels have been set. Firms in the tradables industries do not have this option, at least not to the same extent. As a simplifying assumption, firms in the tradables sector can only make contingent plans, based on given output prices.

This analysis identified a structural wage coordination problem. For example, in situations with practically full employment, firms in the non-tradables sector might be tempted to “bid up” wages to attract workers, which would erode the international cost-competitiveness and profitability of the tradables sector, which was seen as problematic in an era where the economy needed to export to pay for imports and to attract foreign capital. A coordination mechanism that became associated with the Main-course model was to implement a special form of norm based wage setting in the sectors of the economy. The theoretical wage norm was defined by the required rate of return needed to attract capital to the tradables industry. In that way, the wage level could be made to hang together with the evolution of the main-course, as a tendency and over the longer term. The insight was that the wage level would in some periods drift above the main-course norm, in other periods it would sink below.

To illustrate his idea, Aukrust (1977) included a picture showing a line graph labelled “Main course”, increasing as a function of time, inside “The Wage Corridor” indicated by dashed lines labelled “Upper boundary” and “Lower boundary”. As an operational definition of the main-course, the kroner value of average labour productivity was suggested, as long-term stability of the functional income distribution in manufacturing would then follow. While the wage scope was a central concept in the model, the rate of unemployment was not seen as a wage determining factor of any particular importance. Hence, the main-course model can be seen as an example of the ‘wage-exogeneity’ theories of the 1960s that we mentioned in the Introduction.⁹

The model formulated by Aukrust and his colleagues was however a conceptual framework, a theory. It was not a protocol about how to fix the real world wage levels and wage relativities. The actual wage setting process in Norway continued to be based on voluntarism, and on a gradual modernization of the wage negotiation order that the Master agreement of 1935 first had first set in place.

2.4 New challenges for wage-fixing

The organisational landscape was relatively easy to oversee in the first post-war decades. There was only two peak organisations to reckon with, and the union density hovered around 55 %, Stokke (2000). Since then, the union density has decreased moderately, and is 50 percent at the time of writing. Interestingly, there has been an opposite development among employers. The degree of organisation increased from 50 percent in 1965, to 70 percent in 2022, Nymoen and Svarstad (2025, Ch.3).

One consequence of the development is that the number of national organisations increased. Since the mid 1980s, it is custom to reckon with four main peak organisations on the worker side, LO is still the largest, but it is no longer unchallenged by the other peak organisation, in particular when it comes to wage policies and priorities before and during the annual wage settlements, see Nymoen and Svarstad (2025, Ch.5).

⁹Years later, the main-course theory was formulated as an hypothesis about cointegration between the log of the wage level and the log of the main-course variable. Early studies for Norway includes Nymoen (1989,1991), Rødseth and Holden (1990) and Johansen (1995a) among others.

Moreover, if we only consider organisational percentages, we may underestimate the wider set of challenges that economic development and societal changes presented to wage coordination by voluntary collective action. Already in the 1970s, and increasingly in the 1980s and 1990s, the economy became more complex and the labour market organisations more fragmented. These developments, and the challenges they represented for rational *wage-fixing* – to borrow a term from the Meade’s book volume from 1982 – presented themselves in all Western economies. In addition, Norwegian industrialists, politicians and trade union leaders also soon faced the problem of how the super-profitable petroleum producing sector could be taken on board, without breaking down wage norms and formal institutions of wage formation.

It is not too far fetched to say that wage coordination came close to becoming defunct on more than one occasion during the second half of the 20th century. According to historians Thomassen and Øksendal (2017), free collective wage formation became disliked by economists inside and close to the central government administration during the 1970s, who concluded that it had become impossible to use wage setting as an instrument for the policy targets of the era, (Thomassen and Øksendal, 2017, p. 33-35).

In 1973, an initiative to replace annual free collective wage bargaining with a norm decided by a central commission was almost made government policy, but was eventually withdrawn. The 1980s came with wider frustration, which included the organisations, indicating that the peak level oriented system had run its course because wage drift at lower levels undermined wage moderation, e.g., Ibsen et al. (2024).

However, the inherited system continued, aided by intermittent checks. For example in both 1978-79 and 1988-89, normal wage formation was replaced by wage laws, Bowitz and Cappelen (2001). Then, a change occurred after the Norwegian banking crisis. The ‘Solidarity alternative’ of 1992 represented a social pact that revitalized collective bargaining with a pledge to retrain total wage growth, (Thomassen and Øksendal, 2017, Chapter 7.8), Bergh (2009, Chapter 11). It is believable that the success of the ‘Solidarity alternative’ was important for the emergence of the current system of coordination through pattern wage bargaining.

2.5 Pattern wage bargaining

Eventually, the modernization of the economy and of working life did not mean the end of coordination in wage setting. Instead, it has adapted and now operates through pattern wage bargaining, PWB, where manufacturing, as a representative of the tradables sector, concludes the first agreement in the annual rounds of wage adjustments. This determines a norm for wage increases for other sectors to follow.¹⁰ Hence it appears that the manufacturing sector still plays a special role in the data generation process of the national wage level.¹¹

This system of coordination has been favorably assessed in international policy evaluations. In 2019, OECD noted that “the system of collective bargaining based on coordinated annual wage increases works well, providing top-level guidance on wage increases that is anchored in macroeconomic realities” (OECD, 2019b, p.37).

¹⁰Calmfors (2025) provides insightful discussion of pattern wage bargaining in Denmark, Finland, Norway and Sweden.

¹¹It goes beyond the scope of this paper, but a similar characterization may apply to more countries than often thought, cf. Müller et al. (2018).

A significant development of the system has been that following the frontrunner agreement, NHO, “in common understanding with” LO, determines the wage scope (*rammen*) for overall wage growth in manufacturing. This practice stems from the 2013 white paper NOU (2013), in which all members of a broadly representative commission agreed that the scope should be credible and serve as a norm for subsequent settlements. It is therefore reasonable to think of the *ramme* as an intermediate target for national wage growth.

3 A simple formal framework

Based on the above, we investigate the hypothesis that the manufacturing sector’s profitability and ability to pay, the main-course variable, has been the most important trend determining factor of the national wage level over a sample that starts early in the 20th century and ends in 2024.

The framework used for the empirical investigation needs to be allow for the facts that secular changes in preferences, institutions and market have influenced how the wage level has developed over time. In a long historical perspective, reduction of the length of the working day has also been a goal for trade unions.

We begin with the following equation, meant to represent a long-run relationship of hourly wage compensation in the manufacturing sector, w^m :

$$w^m = \alpha_1 mc + \alpha_2 h^n + x, \alpha_1 > 0, \alpha_2 \geq 0, . \quad (1)$$

Since the equation is linear in parameters, we think of w^m as measured in logarithms. On the right hand side, mc is the main-course variable defined as:

$$mc = p^m + z^m, \quad (2)$$

where p^m is the product price on manufactured goods (in practice the deflator of value added at basic values), z^m is value added per hour worked, i.e., the operational definition of labour productivity in manufacturing.

The second explanatory variable, h^n , is normal hours worked per year (in logs), meaning that in the data it will capture not only the development of the normal length of the working day, but also longer holidays and the change from a six to five day working week.

The last variable in (1), x , represents labour market tightness, but also a constant and other deterministic terms.

In the empirical modelling below, we work with a time series for wage compensation per year. Instead of (1) we therefore continue with:

$$w^{am} = \alpha_1 mc + \alpha_2 h^n + h^m + x, \alpha_1 > 0, \alpha_2 \geq 0 \quad (3)$$

where w^{am} is the wage per year in the manufacturing sector, and h^m represents actual hours worked per year in that industry.

We next introduce a second hypothesis: If pattern-wage bargaining has been a feature of the wage setting system, we expect to see stability in relative wages. In the simplest form:

$$w^{af} = w^{am}, \quad (4)$$

where w^{af} is the log of the wage level in wage-following sectors, which we for simplicity think of as the rest of the economy.

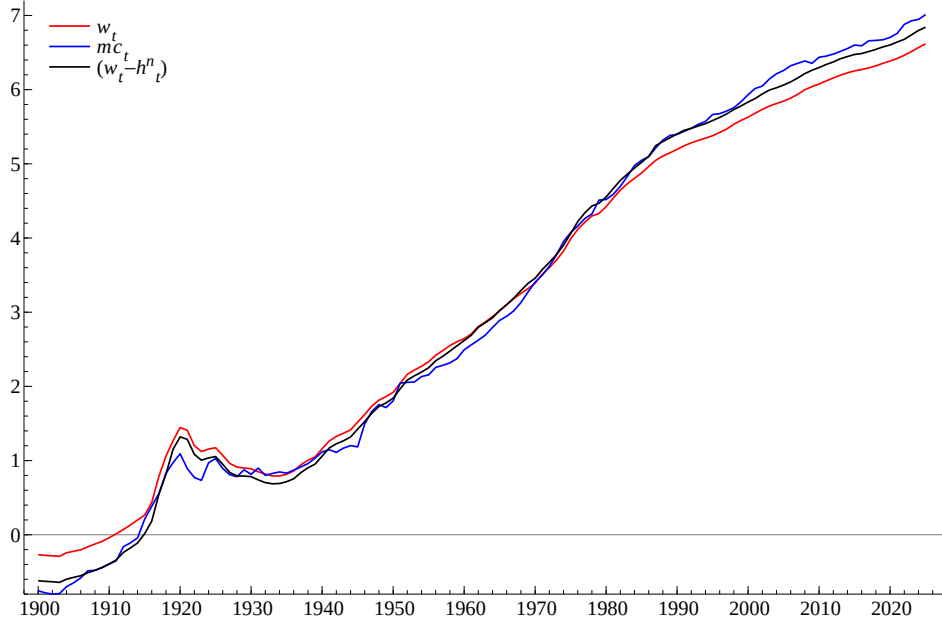


Figure 1: Time plots of the national wage and the main-course 1900-2025. The line graphs have been matched by means of the series.

We can therefore define the log of the national wage level, denoted by w , as:

$$w = \gamma w^{af} + (1 - \gamma) w^{am}, \quad 0 < \gamma < 1. \quad (5)$$

Insertion of the previous equations give:

$$w = \alpha_1 mc + \alpha_2 h^n + h^m + \alpha_3 x. \quad (6)$$

Finally, and again with the reference to the long time perspective, we assume that the number of hours worked per year is proportional to normal hours, denoted by h^n .

$$h^m = \iota h^n, \quad 0 < \iota \leq 1. \quad (7)$$

The above formulations imply the following equation for the national annual wage level:

$$w = \alpha_1 mc + (\alpha_2 + \iota) h^n + x \quad (8)$$

Below, we use econometric methods to investigate the hypothesis that (8) represents a valid long-term relationship. However, as a background, Figure 1 highlights the trend properties of the yearly wage in Norway and the main-course time series in the data set.¹²

The line graph labeled w_t shows the natural logarithm of the national annual wage in Norway 1900-2025. The dashed line labeled mc_t shows the logarithm of the main-course variable defined by the manufacturing sector. For better visualization of the degrees of correlation, the line graphs have been matched by sample means.

The figure indicates a common trend in the w_t data and the main-course time series mc_t . Moreover, they are also correlated in the sub-sample marked by nominal wage reduction in the 1920s, the period of “hard parity-policy” by Norges Bank.

¹²The data set is included in the appendix.

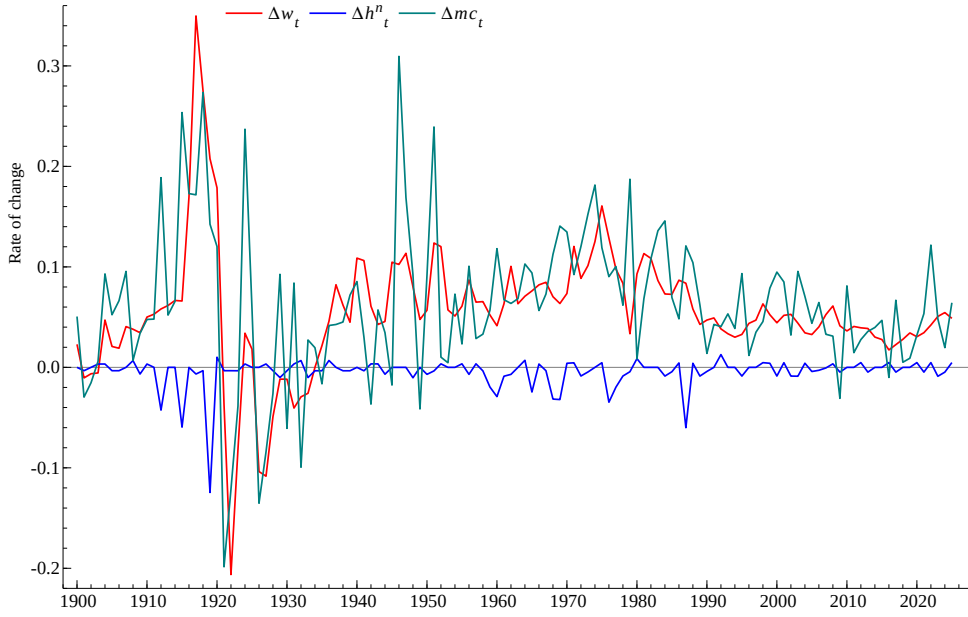


Figure 2: The wage inflation rate, Δw_t , plotted together with the rates of change in the main-course variable, Δmc_t and in normal working hours, Δh_t^n

A rough and ready way to take the secular reduction in normal hours into account, is to subtract h^n from w , which gives the dotted line graph labelled $(w_t - h_t^n)$. Clearly also this adjusted wage series is strongly associated with the main-course variable.

It might be noted that using $(w_t - h_t^n)$ as the wage variable to analyse can be rationalized in two ways. First, judging from previous econometric modelling of manufacturing sector wage formation in times of working time reforms, unions have struggled to maintain the compensation of the hourly wage in the long term, see Nymoen (1989b). The translates to setting $\alpha_2 \approx 0$. Perhaps less debatable, in the long time perspective, is to set $\iota = 1$; the actual working time is proportional to normal working time in the long run.

Second, it is also true that, when the annual wage is adjusted for normal working time as suggested, we obtain a proxy for a national wage per hour. The improved fit with mc , also a per hour variable, is therefore not very surprising.

Below, we will integrate formal analysis of a (potential) long-run relationship with a model of annual change in the national compensation, denoted Δw_t . As a background, Figure 2 plots Δw_t together with Δmc_t and Δh_t^n .

The graphs suggest that there may be a short-run relationship between the national wage and the manufacturing main-course. However, since the volatility is visibly larger for the Δmc_t series than for Δw_t , we can imagine that a (short-term) regression coefficient for wage growth conditional on the change in the main-course will be less than one. The dotted line graph shows the rate of change in the length of the working day and shows that reductions were most frequent in the first part of the sample. The normal working time, which was 3000 hours per year in 1900, had been reduced to below 2400 hours in 1920.

The secular reduction in working time was very modest over the next four decades.

In 1958, the length of the normal working day was 2272 hours per year. But then a new period of significant hours reduction started, which is also visible in the graph. The development brought normal hours down to 1729 in 1987. In 2024 the number of normal working hours was 1676, meaning that the pace of hours reduction slowed down after 1987.

4 Results of econometric modelling and testing

The graphical analysis in Figure 1 supported the hypothesis about a long-run relationship between the main-course variable of the manufacturing sector and the national wage level. However, care must be taken, since relationships between trending economic time series may appear to exist, even when the series are in fact independent, which has been known since the seminal paper by Granger and Newbold (1974), which in its turn resounded Yule’s 1926 warning about “nonsense correlations” between time series.

Cointegration analysis gives a feasible way to correct testing of hypotheses about long-run relationships. One important premise is that the focused time series variables are dominated by a low frequency unit-root. In that case they can be treated as integrated variables, denoted $I(1)$ in the case where the zero-frequency unit-root is not present in the differenced series.

The statistical tests in Appendix A.1 support that, for the purpose of modelling, the time series w_t , p_t^m and z_t^m can be regarded as $I(1)$ -variables. The main-course variable mc_t is defined as the sum of p_t^m and z_t^m . Therefore, mc_t is taken to be $I(1)$ without further assumptions.

4.1 Empirical long-run relationship

To prepare the ground for the analysis we re-express (8) with time subscripts for the variables, and with an explicit error-term ϵ_t as part of the model equation:

$$w_t = \beta_1 mc_t + \beta_2 h_t^n + \beta_3 u_t + c + \epsilon_t \quad (9)$$

In line with the discussion above, the main hypothesis to test is $\beta_1 = 0$, against the alternative that β_1 is positive, and quite close to one. However, given that institutional and other changes noted above, there is reason to expect that the estimation and test results can depend on which sub-sample the model is estimated for.

In (9), x in (8) has been replaced by the rate of unemployment u_t .¹³ It is meant to capture both changes in labour market tightness, as well as being a indicator of historical variations in the balance of power between the worker and capital owner side. c is the constant term.

As noted immediately above, it can be a reasonable simplification to set $\beta_2 = 1$, and therefore to model the approximated wage per hour variable:

$$wh_t = w_t - h_t^n. \quad (10)$$

Hence, we investigate whether

$$wh_t = \beta_1 mc_t + \beta_3 u_t + c + \epsilon_t \quad (11)$$

¹³In logarithmic scale.

is defensible as a cointegration relationship so that $\epsilon_t \sim I(0)$ in contrast to the $I(1)$ -ness of the wage and the main-course specifically. It can be noted if we take as a premise that that nominal homogeneity, $\alpha_1 = 1$, holds in the equation for manufacturing wages setting (1), and wage formation elsewhere hangs close enough together with manufacturing, we expect also β_1 to be close to +1.

Appendix A.2 contains formal test results, indicating that

$$\hat{w}h_t = 0.97mc_t - 0.15u_t - 7.13 \quad (12)$$

is a valid long-run equation.

The estimated β_1 is 0.97, close to the homogeneity mentioned above. Since u_t is the log of the rate of unemployment, the negative coefficient is the correct sign for that variable.

As a check of robustness, we repeated the estimation, with the annual wage w_t , cf. Appendix A.2. Again, the null hypothesis of no cointegration could be rejected at the 5 % level. The estimated coefficient of mc_t was 0.9, a little lower than in (12), which can be said to be consistent with the impression given by the plots in Figure 1.

As discussed in the appendix, diagnostic testing show departures from the normal distribution which gives the statistical foundation of the cointegration test used. To work around this issue, it is more practicable to use a conditional modelling framework instead of the closed VAR.

4.2 Conditional equilibrium correction wage dynamics

It is well known that cointegration implies equilibrium correction dynamics, and vice versa, finding significant equilibrium correction implies rejection of the null-hypothesis of no relationship. Direct testing and estimation of a cointegration in a well specified conditional wage equation also gives a robustness test of the cointegration results immediately above.

Equation (13) shows an empirical wage equation of the equilibrium correction type, ECM hereafter, with $\Delta wh_t \equiv wh_t - wh_{t-1}$ as the dependent variable.

The adjustment coefficient of the lagged wage, wh_{t-1} , is -0.18 , consistent with cointegration. The t-value (below the coefficient estimate) is -7.8 , which is significant as a formal test of the null-hypothesis of no cointegration judged by the critical values in Ericsson and MacKinnon (2002).¹⁴

$$\begin{aligned} \Delta \hat{w}h_t = & \underbrace{-0.180}_{(-7.82)} wh_{t-1} + \underbrace{0.177}_{(7.75)} mc_{t-1} + \underbrace{0.269}_{(8.03)} \Delta mc_t - \underbrace{0.633}_{(-4.83)} \Delta h_t^n - \underbrace{0.019}_{(-5.12)} u_{t-1} \\ & - \underbrace{0.165}_{(-12.8)} SI_{1916t} + \underbrace{0.145}_{(5.22)} SI_{1920t} + \underbrace{0.056}_{(2.31)} SI_{1921t} - \underbrace{0.078}_{(-3.24)} I_{1922t} - \underbrace{0.049}_{(-2.23)} I_{1927t} \\ & - \underbrace{8.10}_{(0.010)} SI_{1935t} + \underbrace{0.077}_{(3.55)} I_{1945t} + \underbrace{0.039}_{(4.3)} SI_{1992t} - \underbrace{1.30}_{(-7.49)} \end{aligned} \quad (13)$$

Estimation method: OLS. Sample: 1904-2025 (T=122)

$\hat{\sigma}_{100} = 2.1$

$\sigma_{\Delta w} 100 = 6.7$

mean $\Delta w = 6.1$

$F_{ar}(2,106) = 3.8 [0.03]$

$F_{arch}(1,120) = 4.35 [0.04]$

$\chi_n^2(2) = 4.7 [0.09]$

$F_{hetxz}(13,103) = 2.0915 [0.0180]$

$F_{f-form}(2,106) = 2.46 [0.09]$

¹⁴With good margin, as the 1 % critical value is -3.8 .

Conditional on cointegration, the implied long-run coefficient estimate for the main-course variable is 0.98, which is practically identical to the estimated β_1 reported in equation (12) above. The coefficient of Δmc_t is lower, but 0.27, with t-value 8, is still evidence of a short-run relationship, as expected from the discussion of the graphs in Figure 2.

The next variable in the first line of the equation is the relative change in the number of normal working hours, Δh_t^n . The estimated coefficient is -0.63 which implies that historical reduction in working time have tended to be partially compensated by higher hourly wages. This applies to the short-term analysis. In the long-term perspective, the cointegration results are consistent with the interpretation that hourly wage compensation is unconnected with the length of normal working hours, cf. Nymoen (1989b).

The last variable in the first line of the equation is the lagged unemployment rate. The t-value of the estimated (short-run) coefficient is -5.1 . The implied long-run elasticity is -0.11 , which is comparable to the u -coefficient in the long-run equation (12) based on the VAR, showing that the estimated relationship is robust with respect to the method used to analyse cointegration.

Following Johansen and Nielsen (2009), these results can be interpreted as robust OLS estimates. Robustness here refers to the model’s ability to handle location shifts in the conditional mean, without allowing them to distort estimation or inference. If such shifts are ignored, standard OLS or IV estimation can produce misleading results due to non-normal residuals and the presence of outliers.

Equation (13) includes two types of indicators to account for structural shifts:

- Impulse indicators ($I_{\text{year}t}$), which capture single-year shocks;
- Step indicators ($SI_{\text{year}t}$), which model persistent shifts in the level of a variable from a specific year onward.

In order to aid interpretation, we note that a step-dummy is set to $+1$ up to and including the year in the subscript, zero thereafter. Hence, the negative coefficient of the first break indicator, SI_{1916t} , implies a positive “wage-shock” that lasts until 1922 when it has been counteracted by two negative shocks, i.e., after both SI_{1921t} and SI_{1922t} have switched from $+1$ to 0. Historically, these sequence of breaks correspond well with the very high wage inflation during WW-I, and the deflationary consequence of the parity policy of the 1920s noted above. As (13) shows, the method picks up three other negative impulses from this period, 1922 and 1927.

The last indicator from the interwar period is the step dummy of 1935, which implies a positive shift to the wage level and which it is reasonable to associate with the Great Reconciliation mentioned above. From the year of Nazi-occupation, the method we applied only picked up a single impulse, for 1945. Interestingly, there is only one break-indicator from the the post WW-II era: the step indicator for 1992 which implies lower annual wage increases after 1992. A not too far-fetched interpretation is that it captures the more lasting effects of the centralized wage settlements of the early 1990s, which were seen at the time as instrumental for restoration of international wage competitiveness, cf. Nymoen and Rødseth (2003), Calmfors (2025) among others.

The inclusion of the indicators was determined using the Autometrics algorithm in PcGive, an automatic model selection tool that applies indicator saturation methods to robustify estimation—see Doornik (2009).

In brief, Impulse Indicator Saturation (IIS) works by creating a full set of time-specific impulse dummies, one for each year in the sample (i.e., T dummies for T observations, in

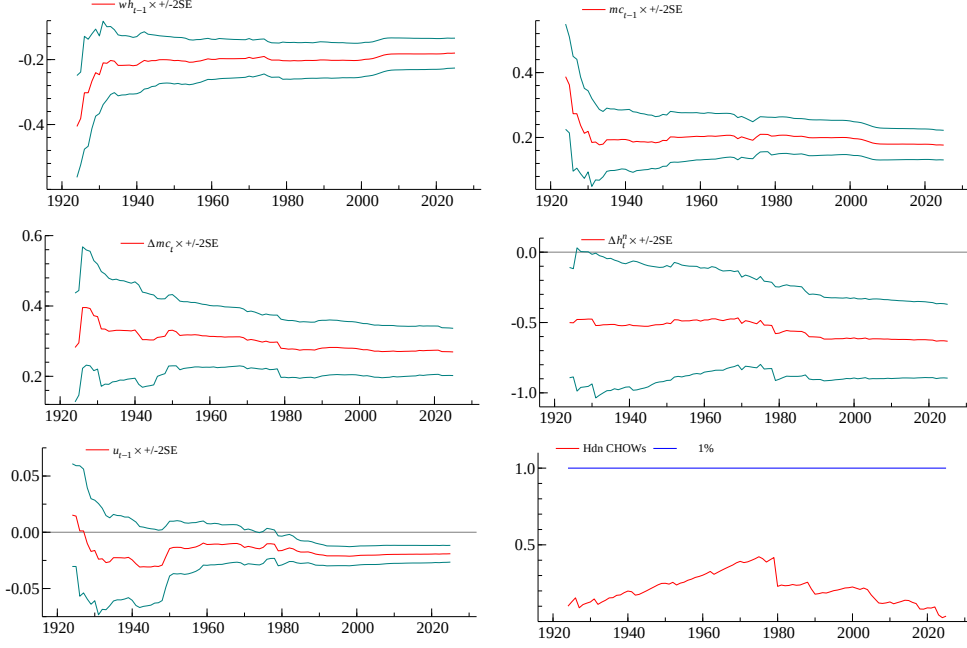


Figure 3: Recursive estimation results: equation (13).

our case 121). This leads to an over-parameterized, or “saturated” model, which cannot be estimated directly due to too many regressors.

To resolve this, a feasible split-sample selection algorithm—developed in Hendry et al. (2008) and Johansen and Nielsen (2009)—is used to reduce the model in a statistically valid way. The goal is to retain only those indicators that capture genuine structural breaks, while discarding spurious ones.

When persistent breaks are expected, Step Indicator Saturation (SIS) is more appropriate. In this case, both IIS and SIS were applied simultaneously. This dual application helps capture both sharp one-off shocks and longer-term regime shifts in the data-generating process.

To focus on identifying only the most large and consequential breaks, we used a stringent target significance level of 0.001. This low threshold reduces the probability of Type-I errors (false positives), effectively prioritizing the detection of large, consequential breaks, even at the cost of potentially missing smaller ones.

Finally, standard residual diagnostic tests—reported below equation (13) show that, considering the long sample period and the volatility of the data, the resulting model has acceptable residual diagnostics.¹⁵

Returning to the economic explanatory variables in equation (13), Figure 3 shows recursive coefficient estimates. The line graphs start in 1924, reflecting that the initial estimation sample is 1904-1923. The panels in the first row illustrate that as the end-of-sample is increased towards 2025, the recursively estimated coefficients of wh_{t-1} and

¹⁵Residual mis-specification tests:

F_{ar} :residual autocorrelation, Godfrey (1978); F_{arch} :autoregressive conditional heteroscedasticity, Engle (1982)
 $\chi_n^2(2)$:non-normality,Doornik and Hansen (2008); F_{hetxz} :heteroscedasticity, White (1980)

F_{f-form} :functional form, Ramsey (1969)

Test statistic values are reported with p-values in hard brackets.

mc_{t-1} in particular stabilize around their full-sample values. The same can be said about about the coefficients of Δmc_t and Δh_t^n in the second row.

The coefficient of the unemployment rate (third row) seems to require the longest estimation samples before it stabilizes. The sign of the estimated coefficient is consistently negative though. Finally, the last graph shows the sequence of break-point Chow-tests (scaled on the 1 % critical values of the F-distribution). Hence, conditional on the eight indicators discussed above (the last one for 1993). the degree of recursive stability can be noted as sign of relative structure.

As noted, the estimated cointegration relationship implied by (13) is:

$$wh_t^* = \hat{c} + mc_t - 0.11u_t, \quad (14)$$

where the “target” variable wh_t^* is defined by the linear combination of variables that is cointegrated with wh_t .

In passing, it is interesting to note that the elasticity with respect to unemployment of -0.11 is close to earlier “wage curve” estimates that used time series that started in the mid 1960s.¹⁶

Consistent with cointegration, we can define the equilibrium correction mechanism:

$$ecm_{wt} = wh_t - mc_t + 0.11u_t, \quad (15)$$

and use it as an explanatory variable in the equilibrium correction equation for Δwh_t . The full sample estimation result for this variant of equation (13) gives:

1904–2025.

$$\widehat{\Delta wh_t} = -0.17 ecm_{wt-1} + 0.26 \Delta mc_t - 0.61 \Delta h_t^n - 1.23 \quad (16)$$

$\begin{matrix} (-8.5) & (8.0) & (4.69) & (-8.25) \end{matrix}$

where we have omitted the break indicators to save notation.¹⁷ Another check of the stability of the coefficients of the economic variables, is to split the sample in a pre World War-II sample and a post-war sample. The two equations are shown in equation (17) and equation (18):

1904–1939.

$$\widehat{\Delta wh_t} = -0.20 ecm_{wt-1} + 0.32 \Delta mc_t - 0.48 \Delta h_t^n - 1.43 \quad (17)$$

$\begin{matrix} (-5.52) & (5.35) & (-2.4) & (-5.36) \end{matrix}$

1946–2024 :

$$\widehat{\Delta wh_t} = -0.14 ecm_{wt-1} + 0.22 \Delta mc_t - 0.81 \Delta h_t^n - 1.02 \quad (18)$$

$\begin{matrix} (-5.41) & (5.17) & (-4.14) & (-5.24) \end{matrix}$

The two sub-sample equations have common features: Significant adjustment of the national wage with the respect to deviations from the secular value of productivity trend defined by the main-course variable. At the same time, the short-run elasticities of the main-course are significantly less than one. The implication is that medium-term wage

¹⁶Nymoen and Rødseth (2003): -0.18 , and Forslund et al. (2008): -0.10 , sample 1965-1997.

¹⁷For completeness, the indicator part of the equation is:

$$\begin{aligned} & -0.17SI_{1916t} + 0.16SI_{1920t} + 0.05SI_{1921t} - 0.08I_{1922t} \\ & -0.05I_{1927t} - 0.07SI_{1935t} + 0.08SI_{1945t} + 0.08SI_{1992t}, \end{aligned}$$

departure from the productivity trend have been a 'normal' phenomenon, and without any permanent de-coupling of wages from productivity.

Figure 4 shows dynamic simulation of model equation (16) over the period 1904-2025. The simulation is conditional on the actual values of the main-course, the unemployment rate and normal hours, as well as the break-indicator (hence, technically they are dynamic ex-post forecasts).

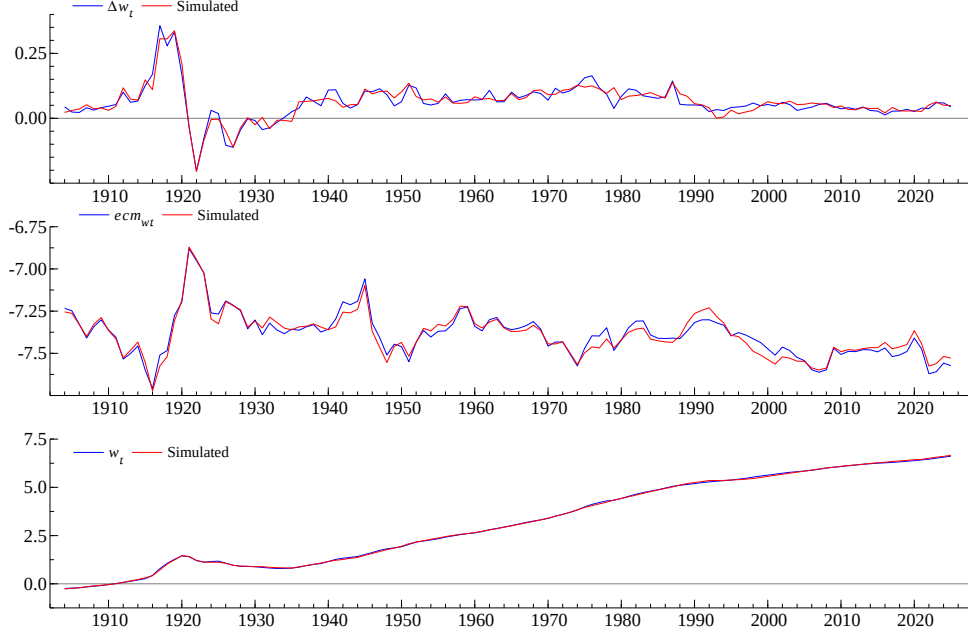


Figure 4: Dynamic simulation of (15)-(16) 1904-2025, conditional on actual values of the explanatory variables (main-course, rate of unemployment, working time and break-indicators).

The line graph in Panel 1 shows actual wage inflation Δw_t , and the dashed line shows the simulated change in the national wage. Panel 2 shows simulated and actual values of the disequilibrium variable ecm_{wt} . The simulated fit is good for both variables, notably also in the relatively long periods where there no are break-indicators included in the estimated wage equation 13.

For completeness, Panel 3 shows the simulated (log) of the national wage level. Clearly, the simulation fits the secular trend in the Norwegian wage level very well, and in the model the variable with enough force to explain the fit is the main-course variable, cf. Figure 1.

4.3 The main-course and purchasing power

Using the results obtained so far, we can define a relationship for the target consumer real wage, rw_t^* :

$$rw_t^* = \hat{c} + z_t^m + (p^m - cpi)_t + h_t^n - 0.11u_t \quad (19)$$

where cpi_t denotes the log of the consumer price index and hence $rw_t^* = w_t^* - cpi_t$.

Panel 1 in Figure 5 plots the actual consumer real wage, $rw = w - cpi$, together with the target real wage rw_t^* . The graphs show that the two series hang together credibly in the long term. Panel 2, 3 and 4 plot rw^* together with z_t^m , the productivity part of the main-course variable, and each of the remaining explanatory variables in (19): the relative price $(p^m - cpi)$, the unemployment rate, u and normal hours.

The overall picture is clear enough: It is the productivity variable z_t^m that explains the secular rise in the purchasing power of the national annual wage. Each of the other variables can be seen as “modifiers” of the z_t^m -graph – they do not represent new or separate trends in the purchasing power of the national wage.

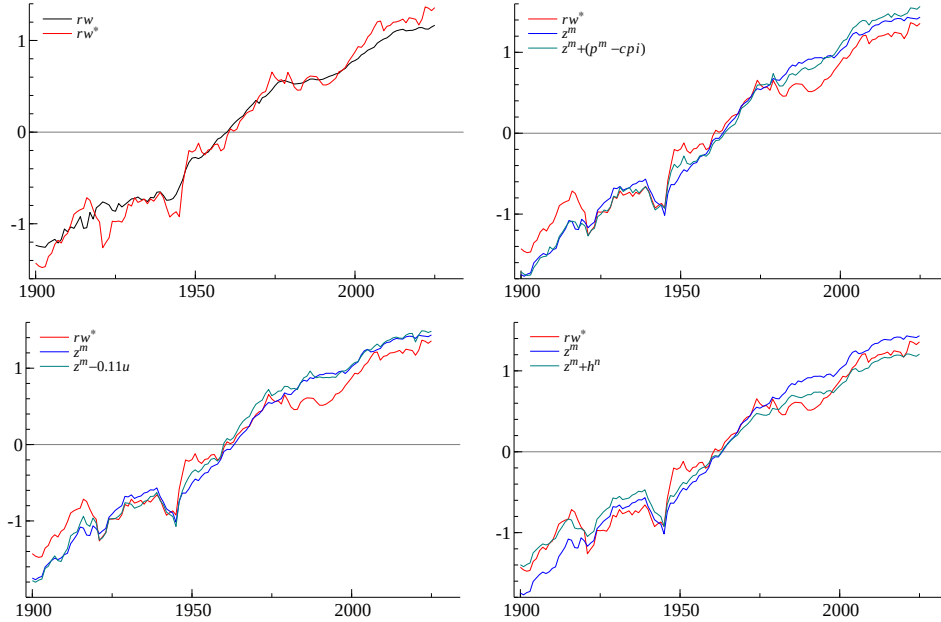


Figure 5: The purchasing power of the annual national wage, rw (actual) and fitted trend rw^* , and the explanatory factors in equation (19). The graphs show demeaned data (deviations from sample averages).

If we supplement the framework in Section 3 with an equation that represents normal mark-up pricing in the wage following sector, and define cpi as a weighted average of the two product prices, the secular growth in $(p^m - cpi)$ should depend on the difference between the two industries' productivities. On the face of it, that difference has not been large enough in the data to create a widening gap between manufacturing productivity and the purchasing power of the annual wage. However, understanding the long-term data generating process of the general price level needs more modelling. It is an interesting challenge for further work.

Table 1: Exogeneity tests.

A. Weak exogeneity tests		
t-values of ecm_{wt-1}		
in wh -equation: (16),	in mc -equation: (22),	in u -equation: (23)
-8.5**	0.14	0.63
B. Invariance tests		
Wage equation exclusion tests		
mc -breaks:	F(7,103) = 1.94 [0.07]	
u -breaks:	F(11,99) = 1.17 [0.33]	
Note: A: ** = significant at 1 %. B: p-values in brackets		

4.4 Exogeneity and invariance

As usual in econometrics, the interpretation of the estimated model becomes more precise if we investigate empirically the exogeneity status of the explanatory variables. The issue is of particular importance for the analysis immediately above, about manufacturing productivity as the main driver of the trend in purchasing power. If there is evidence of reverse causality, meaning that wages drive productivity, the interpretation of the above results must be reconsidered.

There are several dimensions of exogeneity to investigate, Engle et al. (1983), Hendry (1995, Ch. 5). The first is weak exogeneity, which is the condition for valid and efficient estimation,

In a model where I(1) variables are combined to give a linear combination which is I(0), the issue is whether lack of exogeneity of mc_t and u_t implies inefficient estimation of the cointegration parameters (for example β_1 in (9) above).

We have established, in (13) and (16), that ecm_{wt-1} is significant in the conditional wage equation.

A feasible test of the weak exogeneity of mc_t and u_t with respect to the estimation of the long-run relationship, is to test whether ecm_{wt-1} is a significant variable in one or both of the marginal model equations. In Appendix B, the two marginal equations are reported, as obtained by the same method we used to specify the econometric wage equation.

Part A of Table 1 summarizes the results, showing that ecm_{wt-1} is insignificant in the marginal equations (22) and (23). Hence, both mc_t and u_t are interpreted as weakly exogenous with respect to the long-term relationship. The result for mc may be interpreted as an indication of absence of (strong) efficiency wage effects in our data set.

The two marginal equations in appendix, (22) and (23), also give a basis for testing super exogeneity of mc_t and/or u_t with respect to the parameters of (13). Super exogeneity is the joint property of weak exogeneity and invariance of parameters with respect to interventions (structural breaks) elsewhere in the system of variables (their shared data generating process), Engle and Hendry (1993).

A necessary condition for having a test of absence of super exogeneity with statistical power is that there are structural breaks in the marginal model equations, and that the indicator variables that represent the breaks that are unique (not the same as the ones already included in (16)). As documented in Appendix B, this condition is fulfilled in our case.

Part B of Table 1 summarizes the results of the invariance tests, which is reported in

more details in the appendix There are seven unique break-indicators in the mc -equations. Jointly, they are insignificant at the 5 % level when tested for exclusion from the wage equation. Regarding, the breaks in the u -equation there are 11 in in total, and as the table shows they are jointly insignificant in the wage equation.

4.5 Robustness checks

So far, the results have supported the existence of a long-term relationship between the main-course and the national wage level, and that the relationship has been upheld by adjustments of wages and not in a significant way by adjustments of the rate of unemployment or of the main-course variable itself. Moreover, the dynamic wage equation has parameters with a surprisingly high degree of constancy with respect to the detectable breaks in the equations for the two explanatory variables. Another term that describes these results is that the relationship is robust.

In this sub-section, we briefly investigate robustness in a few other directions. In Table 2, the results in the column labelled M0 is the baseline model, i.e., equation (16) with ecm_{wt-1} defined in equation (15). In order to save space the results for the indicator variables are not reported in the table (their role was discussed above).

Comparison of the results for M0 with the results for M1 in the table shows that coefficient estimates of ecm_{wt-1} and Δmc_t are robust with respect to estimation method: OLS in the M0-column and Instrumental Variable (IV) estimation in the M1-column. Concretely, in M1 Δmc_t is treated as an endogenous explanatory variable, with Δmc_{t-1} and the indicator variables in equation (22) as instrumental variables. This means that there are 4 over-identifying instrumental variables. The validity of the over identifying instruments, can therefore be tested by the specification test of Sargan (1958,1964), which is reported in the M1-column and is insignificant at the 5 % level.

It is of interest to test robustness with respect explanatory variables that are frequently used in wage modelling. Cost-of-living is arguably first among such variables.

Table 2: Robustness check of the dynamic equation between the national wage and the main-course. Sample 1904-2024.

Model:	M0	M1	M2
	Baseline, (16)	IV estimation	Cost-of-living
ecm_{wt-1}	-0.17**	-0.18**	-0.14**
Δmc_t	0.26**	0.34**	0.26**
Δh_t^n	-0.62**	-0.55**	-0.52**
Δcpi_{t-1}			0.15**
Sargan-test		17[0.02]	
** = significant at 1 %. For Sargan-test, p-value in brackets.			

Column M2 in the table extends the baseline model with the lagged inflation rate Δcpi_{t-1} . As the result show, the cost-of-living variable is significant and the two incumbent explanatory variables retain their significance, even if the estimates are somewhat reduced compared to the baseline model (M0).

5 Summary and discussion

In this paper, we have investigated the relationship between the Norwegian version of the wage-scope variable—the main-course—and the national wage level over a long historical period. Using impulse and step indicator saturation, we obtained robust coefficient estimates and tested their constancy in light of potential structural breaks, particularly those affecting the main-course variable.

Given the profound changes in economic structure, technology, and labour market participation over time, it would be naïve to expect a fully stable model equation. Our approach instead began with a general model, saturated with indicators capturing structural breaks. Using the machine learning algorithm *Autometrics*, we identified and retained only the most significant breaks in the empirical relationship between the main-course and the national wage.

Interestingly, relatively few break indicators were retained by the estimation method. Those that were can be interpreted as capturing the volatility evident during the first World War, when the Norwegian economy boomed, and during the extended period of stagnation and crises throughout the 1920s and 1930s. This era also marked the formation of the institutional framework for collective bargaining in Norway—based on voluntarism and free negotiation.

From the period of institutionalized collective wage bargaining onward, only two break-indicators were retained. One is an impulse dummy that pick up the effect of a wage ‘freeze’ in 1979. The other is a step-dummy which captures the more persistent effects of the centralized wage settlements of the early 1990s.

Conditioned on the break indicators, the results document a relatively high degree of constancy in the parameters of the relationship. Since the early 1980s, employment in the manufacturing sector—which defines the main-course both in our model and in actual wage formation—has declined significantly. Today, the sector accounts for less than 10 percent of total hours worked in Mainland Norway. It would not be surprising if such a secular shift weakened the legitimacy of the main-course as an anchor for wage growth in the broader economy. However, if such erosion had progressed far, we would expect to observe a breakdown in the empirical relationship—yet we do not.

In terms of econometrics, our findings support the interpretation of the main-course as a super exogenous variable in the wage equation, exhibiting a degree of invariance to structural breaks over the period studied. This implies that the main-course has been a powerful driver of wage levels, extending well beyond the boundaries of the manufacturing sector. This is consistent with the current institutional reality: in the 2020s, the manufacturing sector’s collective agreement continues to serve as the “front-runner” in Norway’s system of pattern wage bargaining.

The robustness of this empirical relationship may reflect the robustness of Norwegian labour market institutions. But institutional robustness is not about rigidity; it is about adaptability to economic, political, and societal changes. So far, institutional adaptation appears to have reinforced, rather than undermined, the leading role of the main-course in national wage formation.

The empirical wage equation we estimate has implications for the development of the purchasing power of the annual national wage. Concretely, the results imply that the singular long-term driver of purchasing power has been average labour productivity per hour in the manufacturing sector. According to our results, three other variables have held force in the shorter term perspective: Reforms in normal hours, the rate of

unemployment, and the relativity between the price of manufacturing products and the consumer price index.

The results obtained in the paper may rekindle discussions about the institutional dependency of the empirical relationship between trend productivity and the secular rise in real wages. Our results are influenced by data from periods with different collective strength (on both sides of the labour market), which suggests that the dependency is a matter of degree, rather than of full dependency or nothing.

In 1980, during an era of industrial unrest in the UK, and based on earlier groundbreaking research in historical economic statistics, and work for the Donovan commission in 1966,¹⁸ Henry Phelps Brown thought that a good question was:

[..] how much had the unions brought their members that productivity would not have brought them in any case?¹⁹

A similar reflection is relevant after our analysis, since we tend to interpret the results as being in agreement with the belief that without collective action, the outcome would have been different. One reconciliation of views may build on the observation that it is not only the existence of a relationship between wage and productivity that is of importance, and which may depend on institutions: It is also the strength. In the long-term perspective, it does matter for the functional income distribution whether the wage elasticity with respect to productivity is 0.5 or 0.75, or indeed unity, as we found as a data acceptable restriction on the relationship.

Our empirical results also lead to new research questions. For example, the roles played by productivity per hour and the length of the normal working year in the decomposition of annual wage should be researched in more detail. In the recent research literature of long historical trends in real income, more attention has been given to working time. Taking working time into account has led to drastic revisions of economic historians' conceptions of real wages during the Industrial Revolution, cf. Humphries and Weisdorf (2019) and the references therein.

To further analyse wage dynamics over the long term, future research will also need to move beyond the reduced-form approach taken here and towards multiple equation modelling. Such an approach may also give better insight into the mechanism that underlie the robustness of the wage-productivity nexus documented by analysis of the aggregated time series.

A Integration and cointegration analysis

A.1 Unit-root tests

Table 3 contains results for Augmented Dickley Fuller (ADF) tests of unit-roots in the key time series, hence the null hypothesis of the tests is that the data generating processes contain the low-frequency unit-root.

The reliability of the ADF-test depends on the correct specification of the deterministic part of the underlying regression model. In Table 3, the regressions used for the three level series: w_t , p_t^m and z_t^m include both a constant term and a deterministic trend, indicated by Constant+Trend. This secures that the regression models are congruent with

¹⁸Phelps Brown (2012a)

¹⁹Phelps Brown (2012b)

Table 3: t-ADF test statistics. Sample 1903-2025

Constant+Trend			Constant		
w_t	p_t^m	z_t^m	Δw_t	Δp_t^m	Δz_t^m
-2.4	-2.1	-1.4	-3.8**	-7.1**	-9.5**
Memo: ** indicates significance at 1 % critical level. D-lag = 0: $\Delta p_t^m, \Delta z_t^m$; D-lag = 1: p_t^m, z_t^m ; D-lag = 2: Δw_t^m ; D-lag = 3: w_t^m .					

the trending behaviour of the series, both under the null hypothesis and the alternative hypothesis (no unit-root). In the regressions used for the tests of Δw_t , Δp_t^m and Δz_t^m , only a constant term is required, since these differenced series do not contain clear trends.

A second issue is the degree of stochastic augmentation used in the unit-root tests, which refers to the number of lags of the differenced series (D-lag) included in the regressions. If D-lag is set too small, the ADF-test is invalidated by residual autocorrelation. If it is set higher than needed to achieve approximately white-noise residuals, the ADF-test may become inefficient in the finite sample we in practice use (hence inflate Type-2 error probabilities). The chosen of D-lags are reported in the Memo-part of the table. They were based on the significance of the t-values of the lagged differences.

A.2 Cointegration tests

To test and estimate the hypothetical long-run relationship (11), we used the vector autoregressive (VAR) cointegration method of S.Johansen (1995), as implemented in the software package *OxMetrics*, see Doornik and Juselius (2018).

Table 4 shows results for the Trace-test of reduced rank of the long-run matrix of two models, VAR(1) and VAR(2). The number of lags for the endogenous variables is two in VAR(2), and one in VAR(1). Both tests are based on an assumption of weak-exogeneity of u_t with respect to the cointegration parameters.

With two endogenous I(1) variables, the rank can be wither zero (no cointegration) or 1. The table shows that rank = 0 can be rejected at a significance level close to 5 % when VAR

Table 4: Cointegration tests based on VARs with wh_t and mc_t as endogenous, u_t as weakly exogenous and unrestricted constant. VAR(2): Two lags of the endogenous, one lag of u_t . VAR(1): One lag of the endogenous, one lag of weakly exogenous u_t .

VAR(1)			VAR(2)	
1904 - 2025			1904 - 2025	
rank	Trace	[p-value]	Trace stat	[p-value]
0	93.23	[0.00]	30.47	[0.056]
1	10.89	[0.219]	7.34	[0.539]
Trace and p-value incorporate the Bartlett small sample correction, Johansen (2002)				

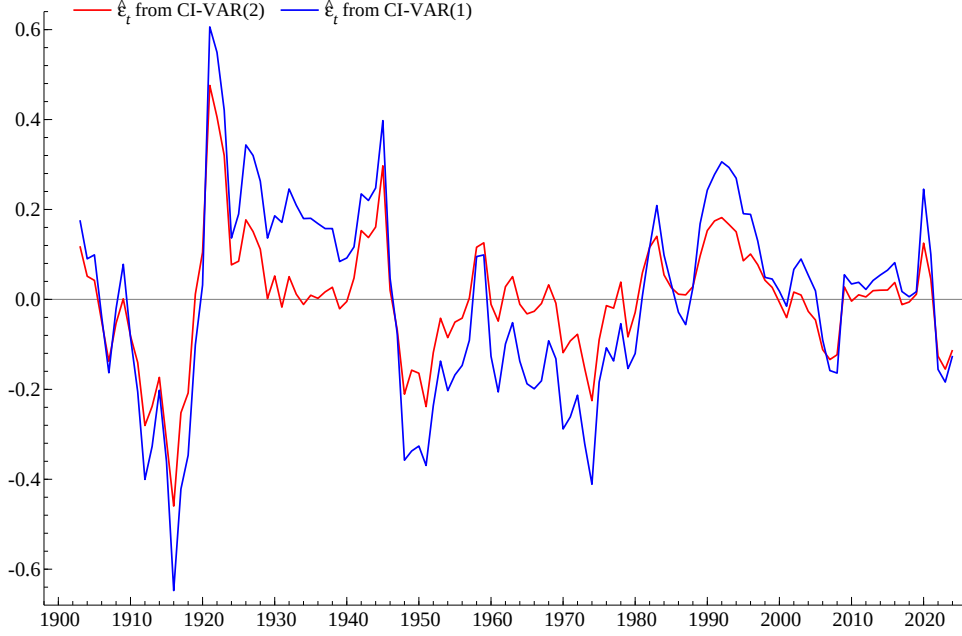


Figure 6: Plots of the residuals defined by the two long-run relationships in (20).

Conditional on rank = 1, the cointegrating relationships of the two models become:

$$\begin{aligned}
 \text{CI-VAR(1)} : wh_t &= 0.961mc_t - 0.31u_t - 7.10 + \hat{\epsilon}_t \\
 &\quad (116.0) \quad (10.4) \\
 \text{CI-VAR(2)} : wh_t &= 0.969mc_t - 0.153u_t - 7.25 + \hat{\epsilon}_t \\
 &\quad (114.9) \quad (4.2)
 \end{aligned} \tag{20}$$

with t-values in brackets below the estimates.²⁰ The coefficients of mc_t are nearly identical and close to long-run homogeneity (+1). There is a more notable difference between the two estimated elasticities with respect to the unemployment rate. However, since the sample variance of u_t is much lower than of mc_t , that difference does not impact qualitatively on the properties of the relationship, see Figure 6.

As noted in the main text, we do not get clear empirical guidance about the treatment of normal working time per year, hn . The main result about cointegration between wage and main-course is however robust. When the analysis is done with w_t instead of wh_t , the tests give rank=1. The estimated cointegration relationships become:

$$\begin{aligned}
 \text{CI-VAR(1)} : w_t &= 0.895mc_t - 0.267u_t + \hat{c} + \hat{\epsilon}_t \\
 &\quad (136.3) \quad (15.8) \\
 \text{CI-VAR(2)} : w_t &= 0.901mc_t - 0.141u_t + \hat{c} + \hat{\epsilon}_t \\
 &\quad (141.8) \quad (5.2)
 \end{aligned} \tag{21}$$

The results are similar to (20), but with somewhat more notable departures from nominal homogeneity.

Although the results of the cointegration analysis are economically meaningful, and robust in the directions that we have investigated, diagnostic tests of the residuals show departures from gaussian properties. Non-normality of VAR-residuals is a contradiction of the assumptions of the statistical framework of the cointegration test.

²⁰The two constant terms are without t-values as they are sample means of the long-run relationships.

A plausible explanation for the diagnostic problem is that there are structural breaks in the sample period that needs to be taken into account. To do this by inclusion of break-indicators in the VAR used to test rank-reduction has implications for the critical values, and is left as a task for further work. Instead, we have used the approach that we document in the main text and make use of a conditional model which has well behaved diagnostics after objective selection of break indicators has been applied.

B Testing exogeneity

To implement the exogeneity tests, marginal model equations were specified for mc and u . For the mc -equation, the 0.1 % target significance level was used.

In (22), the t-value of ecm_{wt} is -0.4 , which does not reject the hypothesis that the main-course variable is weakly exogenous with respect to the parameters of the long-run relationship.

$$\Delta \widehat{mc}_t = \underset{(0.14)}{0.01} ecm_{wt-1} + \underset{(5.51)}{0.37} \Delta mc_t + \underset{(0.25)}{0.08} + 8 \text{ impulse indicators} \quad (22)$$

Estimation method: OLS. Sample: 1904-2025 (T=122)

$$\begin{array}{lll} \hat{\sigma}100 = 4.85 & \sigma_{\Delta mc}100 = 7.5 & \text{mean } \Delta mc100 = 6.4 \\ F_{ar}(2,109) = 1.72 [0.18] & F_{arch}(1,120) = 0.95 [0.33] & \chi_n^2(2) = 3.11 [0.21] \\ F_{hetxz}(4,109) = 0.83937 [0.5031] & F_{f-form}(2,109) = 3.17 [0.05] & \end{array}$$

That it is admissible to restrict the coefficient of ecm_{wt-1} to zero in (22) is also important for the interpretation of the simulated fit in Figure 4: It represents formal support for the theory that it is the main-course that is driving wages, and not the other way around.

There are eight significant break-indicators in (22), for which the detailed estimation results have been omitted to save space.

The invariance of the wage equation with respect to the breaks can be tested by a standard omitted variable test in the wage equation. The test of joint insignificance of seven unique break indicators from the mc -equation ($I_{1915}, I_{1918}, I_{1924}, I_{1926}, I_{1932}, I_{1946}, I_{1951}$) gave as result $F(7, 103) = 1.94[0.07]$.²¹

When we undertook the same type of modelling exercise for the (log of) the rate of unemployment we obtained equation (23):²²

$$\begin{aligned} \Delta \widehat{u}_t = & \underset{(0.63)}{-0.12} ecm_{wt-1} + \underset{(5.81)}{0.35} \Delta u_{t-1} - \underset{(7.96)}{0.39} u_{t-1} + \underset{(0.86)}{1.26} \\ & + 4 \text{ impulse indicators} + 9 \text{ step indicators} \end{aligned} \quad (23)$$

Estimation method: OLS. Sample: 1905-2025 (T=121)

$$\begin{array}{lll} \hat{\sigma}100 = 17.6 & \sigma_u100 = 72.3 & \text{mean } u = 0.93 \\ F_{ar}(2,102) = 1.0 [0.36] & F_{arch}(13,101) = 1.43 [0.16] & \chi_n^2(2) = 1.0 [0.61] \\ F_{hetxz}(13,101) = 1.43 [0.16] & F_{f-form}(2,102) = 3.44 [0.04] & \end{array}$$

²¹One of the six indicators in (22) is I_{1921} , which is perfectly collinear with SI_{1920} the SI_{1921} in the wage equation, this means that the number of unique mc -break indicators are seven.

²²With target level 0.1 % Autometrics selected very few break indicators, and none after 1920. To obtain the reported marginal equation, target level 1 % was used

The statistical insignificance of ecm_{wt-1} is in support of the hypothesis that u_t is weakly exogenous with the respect to the parameters of the long-run relationship for the wage-level. As noted in (23), the marginal model of Δu contains 13 indicator variables (4 impulses and 9 step-indicators). Years with break indicators are from different parts of the sample, for example 1910, 1947, 1949 1957, 1979 – and from both 2020 and 2022.

The result of the exclusion test of eleven unique u-break indicators from the wage equation gave $F(11, 99) = 1.17[0.33]$.²³ Hence, u_t can be regarded as super exogenous with respect of the parameters of the relationship between the wage and the main-course.

C Data

Table 5 displays variable definition and sources. The data series of the variables W, PM, ZM and HN were obtained from a historical time series project commissioned by Statistics Norway which will be published as Harildstad (2026).

Table 5: Variable definitions and sources.

Variable		Unit	Source
W	Wage per year.	1000 kroner	Harildstad (2026)
PM	Price deflator of manufacturing value added, basic value.	2015=1	Harildstad (2026)
ZM	Average labour productivity, manufacturing	Kroner, 2015 prices	Harildstad (2026)
HN	Normal working hours	Hours	Harildstad (2026)
U	Unemployment rate	Percent	Nymoen (2017), Nav
CPI	Consumer price index	2005=1	Nymoen (2017), SSB
SSB: Statistisk sentralbyrå (Statistics Norway)			
Nav: Norwegian Labour and Welfare Administration			

The historical time series of the unemployment rate, U, and of the consumer price index, CPI, are extensions of the series documented in Nymoen (2017), which had 2015 as end of sample.

Unemployment rate.

NAV has published a historical data for registered unemployed in percent of the labour force is available since 1948.²⁴ It represents the longest consistent time series for the Norwegian unemployment rate. We use this series for the period 1948 to 2024.

For the period with 1940-45 (Nazi occupation) and 1946-47, the interpolation method described in Nymoen (2017, Data definitions and sources) has been used. For the interwar period 1920-1939 we use the series from Grytten (2008, Table 3). For 1903-1919 we used a scale adjusted version of the the time series of unemployment rate of trade union members in percent of union membership.²⁵

Consumer price index.

²³u-break indicators I_{1921} and SI_{1945} were perfectly collinear.

²⁴<https://www.nav.no/no/nav-og-samfunn/statistikk/arbeidssokere-og-stillinger-statistikk/historisk-statistikk>

²⁵The data are taken from SSB (1948, Table 187) and SSB (1958, Table 162).

The series documented in Nymoen (2017, Data definitions and sources) has been extended to 2024 by the use of the yearly figures of the official consumer price index²⁶.

Data availability The time series of the variables in Table 5 are available in xlsx-file format upon request.

Definition of the log-transformed variables used in the analysis:

$$\begin{aligned}w &= \ln(W), p^m = \ln(PM), z^m = \ln(ZM), mc = p^m + z^m, h^n = \ln(HN) \\u &= \ln(U), cpi = \ln(CPI).\end{aligned}$$

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²⁶<https://www.ssb.no/statbank/list/kpi>

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