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Saket Hishikar

Effective Hedges against Inflation for Pension Funds

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Effective Hedges against Inflation for Pension Funds

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SAKET HISHIKAR

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University Supervisor(s): Prof Dr. Frank de Jong

Second Reader: Prof Dr. Lieven Baele

Company Supervisor(s): Dr. Pieter van Foreest & Jan Fokkens MSc.

Dutch pension funds that index their pension liabilities to wage inflation commonly use the inflation linked derivatives based on the Harmonized Index of Consumer Prices (HICP) as a best proxy for wage inflation. This thesis provides a justification for this strategy and the assumptions under which this strategy can be adopted. By using the techniques of panel co-integration and vector error correction we establish that wages and prices are co-integrated. The nominal wage adjusts to changes in price by a factor slightly more than one in the long-run. Under the assumptions- (i) that the emergence of HICP as a new measure of inflation does not alter the behaviour of economic agents, and (ii) the long-run co-integration equation assumes a trend, this implies that for pension funds that index their pension liabilities to wage inflation, derivatives based on HICP are effective hedges in the long-run. Presence of co-integration ensures that basis risk is minimized in long-run. We also demonstrate that current ALM framework which models the long-run as a series short period as against the co-integration framework that views long-run as one time block can give very different outcome. The current ALM framework needs modifications to incorporate the long term characteristics of pension funds. We also argue that, the co-integration framework can be useful in finding cost-effective hedges for non-tradable risks like wage inflation.

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Foreword

This master's thesis is an outcome of my four months' internship at the Algemene Pensioen Groep (APG), Amsterdam. This period of four months has been a process of intense learning and application and I take this experience as a lifelong asset. First and foremost, I thank Dr. Pieter van Foreest and Jan Fokkens, my supervisors at APG who put the confidence in me and entrusted me with this research project at APG. This work has benefitted from their suggestions and guidance at all stages.

I owe my sincere thanks to Prof. Dr. Frank de Jong, Tilburg University, who accepted the responsibility as my university supervisor. I thank him for all the valuable insights he has provided to bring this thesis to its just completion.

I also thank Rubin Laros, Rob van den Goorbergh and Bart Kuijpers at APG for helping me at various times during my stay at APG. I extend my special thanks to Vesile Kutlu, my batch mate from Netspar master, for all the help with complex econometrics and Stata.

SAKET HISHIKAR

I

Introduction

Pension payments in the Netherlands are adjusted for inflation to preserve the standard of living of the pensioner. The adjustment for inflation can either be based on price inflation or wage inflation (sector specific or economy wide). This exposes the pension fund liability to inflation risk. It is important that inflation risk is appropriately managed to meet the ambition of inflation compensation on a sustained basis.

The financial literature extensively documents inflation hedging properties of conventional financial assets (equity, bonds, commodities, real estate, and commodity currencies). In this study we digress from these assets and focus on inflation indexed bonds and inflation derivatives (swaps, futures). Inflation linked assets are recent innovations. Pension funds that index pension liability to price inflation benefit from such products directly because inflation linked products provide an efficient mechanism to mitigate the inflation risk on their liability. **However for pension funds that use wage inflation to index pension liabilities, lack of market for wage inflation means that pension fund have to find the best available proxy to hedge this risk. Common practice among the Dutch pension funds is to use the available (price) inflation linked products as the best proxy for wage inflation. In this thesis we provide a justification for this strategy and the conditions/assumptions under which this strategy can be adopted.**

The inflation linked products in Europe are linked to the Harmonised Index of Consumer Prices (HICP) used by the European Central Bank to conduct its monetary policy. Dutch pension funds and their participants are exposed to the Dutch inflation and not to HICP. Secondly, the pension liabilities are typically indexed to (sector specific) wage inflation. Hence there are two mismatches- between HICP and the Dutch CPI, and between wage inflation and HICP, the indicator of Euro area price inflation. We investigate qualitatively and quantitatively the short-run and the long-run relationships between HICP and the Dutch CPI; and wage inflation and HICP. This helps in ascertaining the basis risk that may arise when HICP linked derivatives are used for hedging inflation risk of pension funds.

Traditional macroeconomic theory, however, looks at wages and prices from very different perspective. These variables have been used as tools for macro stabilization policies. Hence the macroeconomic literature on modelling wages and prices has developed from this angle. We follow this strand of literature and use the techniques therein from the perspective of pension funds. The intuition for this approach is that if there is a long-run convergence between the Dutch wage inflation and the price inflation based on Euro area HICP, then inflation derivatives based on Euro area HICP are also effective hedging instruments in the long-run for those pension fund that index their pension liabilities to wage inflation.

In statistics the long-run relationships are modelled using the technique of co-integration. This thesis specifically uses the statistical techniques of panel co-integration and error correction to establish that a long-run relationship between wages and prices exists. We use panel data on wages and prices for the twelve countries that form the HICP index and conclude that the long-run elasticity of nominal wages with respect to changes in prices is greater than one. In the short-run wages and prices can differ substantially. The implications of these findings are as follows:

1. Wages and prices are co-integrated. The nominal wage adjusts to changes in price by a factor slightly more than one in the long-run. That is, the estimated hedge ratio is greater than one. Given the assumptions- (i) that the emergence of HICP as a new measure of inflation does not alter the behaviour of economic agents, and (ii) the long-run co-integration equation assumes a trend, this implies that for pension funds that index their pension liabilities to wage inflation, derivatives based on HICP are effective hedges in the long-run. Hedging is effective because co-integration between wages and prices ensures that basis risk is minimised in the long-run even though wage inflation diverges from price inflation in the short-run.

2. As a corollary to the previous argument it is unadvisable to hedge the risk of wage inflation on a short-run basis using inflation linked instruments, because the hedge quality is poor due to large short-term deviation between wage inflation and price inflation. Hence it is recommended that pension funds hedge the risk of indexation using HICP only on a long-run basis.

3. Current practice in ALM is to view the long-run as a series of short periods. This is typically achieved by generating economic scenarios for one time step in the future till the desired time horizon. This approach therefore assumes that a single long-run decision is equivalent to a series of short-run decisions. However co-integration views long-run as one single time block. This difference can lead to very different outcomes. Furthermore, pension funds are perceived as long-run investors. Hence the current ALM practices should be modified so that they address the long-run decision making characteristics of pension funds.

4. If a pension fund wants to buy a cost efficient hedge for wage-inflation, lack of markets for such a risk implies paying a high premium for the insurance. The co-integration framework as suggested here helps in finding cost effective hedging instruments like inflation linked derivatives for the risk of wage inflation. This is because when wages and prices are co-integrated, the correlation between wage and prices will increase over time and tend to one. Thus the hedge effectiveness of inflation linked derivatives improves over time. The liquid market for HICP derivatives makes the hedge cost efficient.

This thesis proceeds in the following direction. We start with a brief introduction of the key variables wage inflation and price inflation – that is how inflation is measured in the Netherlands and in the Euro area and what is the possible connection between wage inflation and price inflation. Second, we revisit the economic theory on wage-price dynamics. Next, we review the empirical literature on wage-price dynamics. We crystallise the approach to address the problem for pension funds in light of the available literature. The chapter on econometric methodology is somewhat technical and readers may skip this without loss of continuity. We conclude by summarizing the findings and their implications. We also present possible extensions and refinements of this model for future research.

II

Key variables – Wage and Price inflation

Wage and price inflation are the key variables in this study. This chapter is devoted to understanding the measurement of price inflation, its link to wage inflation and the policy of wage inflation compensation by pension funds. This is essential because of two reasons: firstly, the dynamics of wages at the macro level and their translation at fund level will have variance because for many pension funds wage inflation compensation is conditional on the health of the fund. Secondly, there are many measures of inflation (price indices), some specific to the Netherlands and some common across all the members of the European Union. Appreciating the differences between these measures of inflation is important because some, but not all, measures form the underlying of inflation linked derivatives. We first cover price inflation followed by wage inflation.

Measures of price inflation:

There are two principal measures of price inflation in the Netherlands. The consumer price index published by the Central Bureau of Statistics (CBS) and the Harmonised Index of Consumer Prices (HICP) also published by CBS. The former (henceforth referred to as Dutch-CPI) is a measure of inflation indigenous to the Netherlands. The latter is a measure introduced by the European Central Bank (ECB) more suited for international comparisons and is a component of the aggregate HICP[1] used by the ECB for its monetary policy for the Euro area.

The Dutch-CPI reflects the change in the value with reference to a base year of a basket of goods and services bought by an average consumer in the Netherlands. The CPI has been widely used in indexation of commercial contracts, wage negotiations, for the indexation of rents and annuities and to adjust tax tables. **The HICP on the other hand is a special purpose measure applicable to all member states of the European Union (EU)**. The ECB uses a weighted average of HICPs of the 12 member states- Austria, Belgium, Finland, France, Germany, Ireland, Italy, the Netherlands, Portugal, Spain, Greece and Luxembourg, to monitor inflation in the EU. The use of HICP is a move towards standardization of the measurement of inflation across various member states. **The main difference between the Dutch-CPI and the HICP is the composition of the basket of goods on which they are based.** The following table[2] gives a qualitative comparison of Dutch-CPI and HICP in general:

[1] We use HICP-EU to refer to the aggregate index and HICP-Dutch for HICP index specific to the Netherlands.
[2] Form of the table is adapted from Annexure Table 2: Henning and Mariagnese (2005); 'Inflation Measures: Too high-too low-internationally comparable?', OECD Paris, 21-22, June 2005. Changes in the contents made as per requirement.

Table 1
Qualitative comparison of Dutch-CPI and HICP[3]

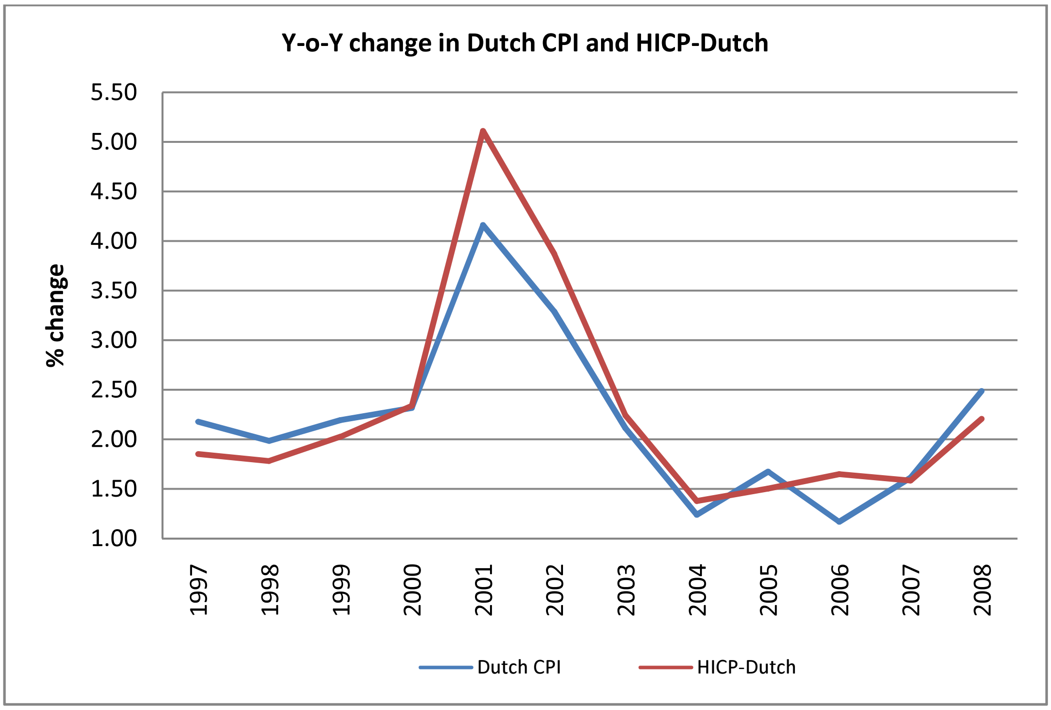
	Harmonised index of consumer prices	Dutch-Consumer price index
Definition	Measure of the average price changes of goods and services available for purchase on the economic territory of the member state for the purpose of directly satisfying consumer needs	Measure of price change for goods and services acquired by households for consumption purposes
Geographic and population coverage	All households on the territory of the member state (includes the private consumption by foreigners in the Netherlands)	Resident population including price increases in private consumption by Dutch people abroad
Item coverage	Private consumption *excluding*: • Owner occupied housing • Consumption-related taxes • Subscription and membership fees for sports clubs, social clubs	Private consumption *including*: • Owner occupied housing • Consumption related taxes • Subscription and membership fees for sports clubs, social clubs • *Excludes* some costs paid for health care
Formula	Laspeyres type	Cost of living index
Weight update interval	Annual review, (at least 5 year)	5 years (plans to move to yearly updates)
Elementary aggregation formula	Ratio of geometric or arithmetic mean	Ratio of geometric or arithmetic mean
Classification	Classification of Individual Consumption According to Purpose (HICP)	Classification of Individual Consumption According to Purpose
Levels of details	94 classes/160 sub-indices	259 sub-indices
Purpose/Objective	Pure inflation measure	Multiple purposes- cost of living, compensation index, reference for domestic policy etc.
Data dissemination	Monthly, not seasonally adjusted	Monthly
Variants	• HICP excluding (ex) energy • HICP ex energy, food, alcohol, & tobacco • HICP ex unprocessed food • HICP ex energy and seasonal good • HICP ex tobacco	• Regular CPI • CPI adjusted for consumer taxes like VAT

[3] Description of HICP in this table is applicable to all the member countries. The weighing of individual items within the basket varies across countries as per taste and preferences of that country. The purpose of this table is to highlight the subtle differences between Dutch CPI and HICP in general.

Weights assigned to items covered within the HICP basket vary across member countries to keep country specific HICP in line with the tastes and preferences of that country. However large variations in weights are regulated and adjustments are made to ensure that final weighted HICP used for monitoring inflation in EU is not adversely affected. Inflation linked derivatives that trade in financial markets use the “HICP excluding tobacco” variant based on the average of twelve countries as its underlying. Hence in financial markets parlances HICP generally refers to HICP excluding tobacco and is denoted as HICPxT. There is no economic rationale for this choice. This choice is an accident in history and is now just an accepted market convention.

From the qualitative comparisons in Table 1 we can ascertain that Dutch-CPI and HICP are broadly similar. The notable differences include the inclusion of imputed rent of ‘owner occupied dwelling’ and coverage of ‘medical cost’ in Dutch-CPI. The difference in purpose of construction should not lead to large variations in calculated inflation from the two measures. Further there is a greater harmonisation in recent years to bring Dutch-CPI in line with HICP-Dutch. Because of these reasons there is a convergence between Dutch-CPI and HICP-Dutch (See Graph 1). The figure presents the yearly series for year-on-year changes in Dutch-CPI and HICP-Dutch from 1997 to 2008.

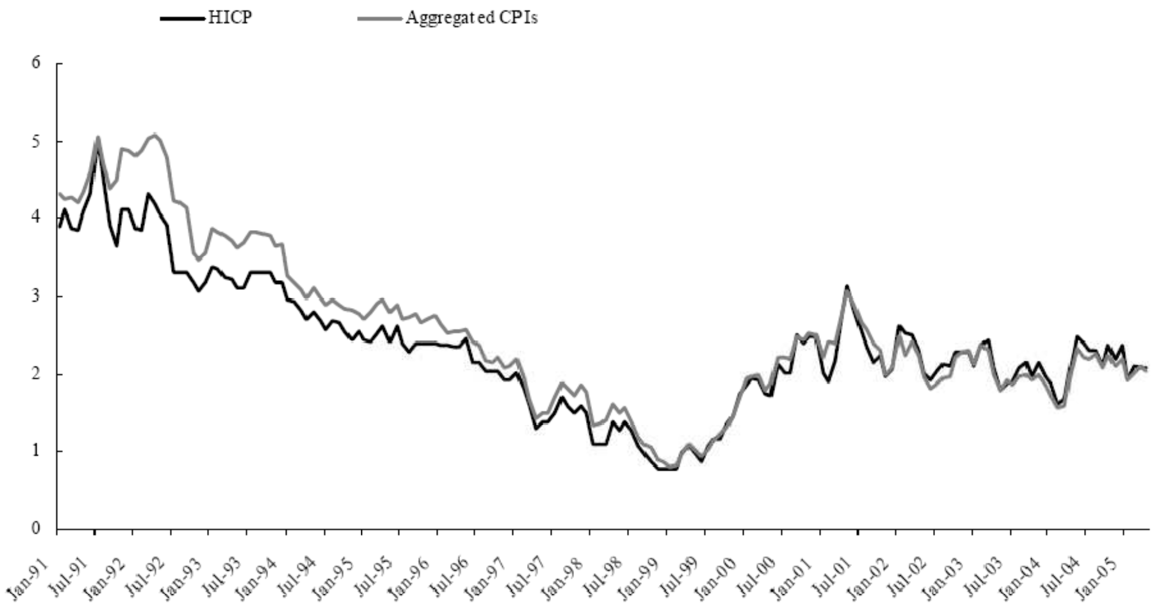
Graph 1



Source: OECD Stats & Eurostat. (Base 2005=100) Correlation: 0.95, Tracking error: 0.39% (STD point). For monthly data on same period the tracking error is 0.42% (STD point)

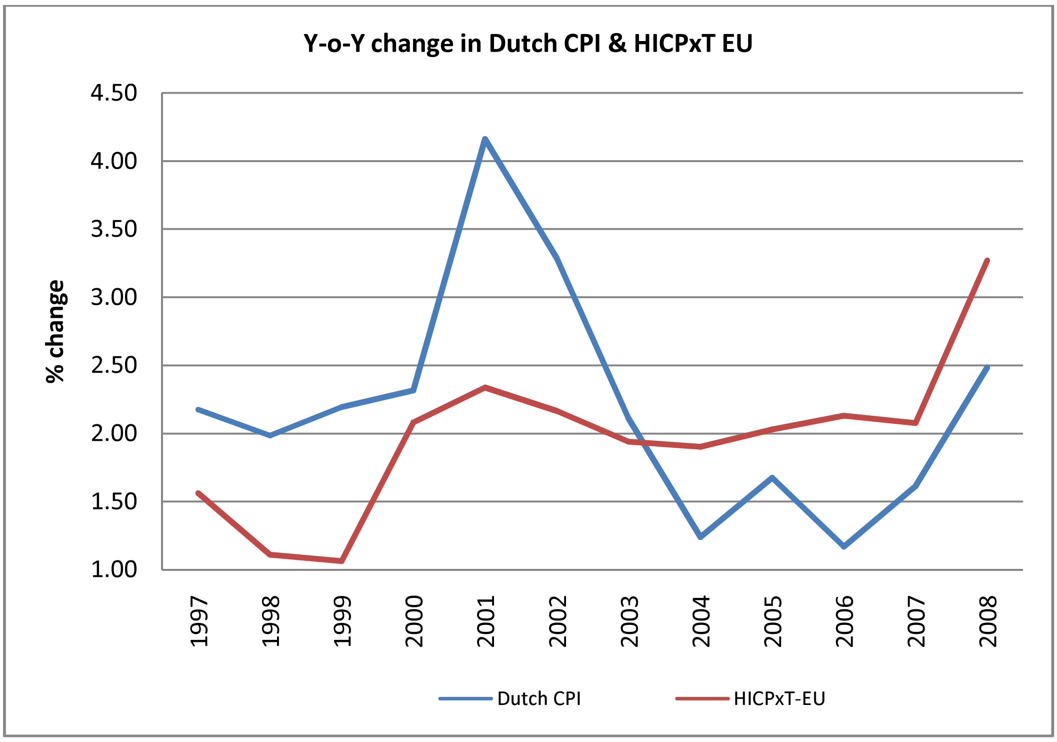
Similar convergence between HICP and CPI has been observed for other member countries. Henning and Mariagnese (2005) highlight this point using the Graph 2 shown below. The importance of convergence will be apparent in the arguments made in Box 2 later.

Graph 2
HICP and aggregate CPIs for the Euro area (annual % change)



Source: Henning and Mariagnese (2005)

Graph 3



Source: OECD Stats, Eurostats. (Base 2005=100) Correlation: 0.25, Tracking error: 0.89% (STD points)

The co-movement between HICPxT EU and Dutch-CPI is less pronounced at the aggregate level. The comparison for the Dutch-CPI and HICPxT EU are presented in Graph 3. **It is at times argued that for pension funds, a domestic price index like the Dutch-CPI is more relevant than an aggregate Euro area price index like HICPxT for hedging. On this point it is worth mentioning that the ECB's quantitative definition of price stability is based on Euro area HICP. The use of HICP is then logical**. Secondly, the use of HICP is further bolstered in light of substantial convergence in actual inflation and inflation expectation within the Euro area[4].

The comparisons presented thus far are for the sake of developing some intuition about the characteristics of domestic measures of inflation and their standardised counterparts in the EU. A more relevant comparison from the point of view of pension funds is between wage inflation and price inflation. Hence wage inflation is the focus in the following paragraphs.

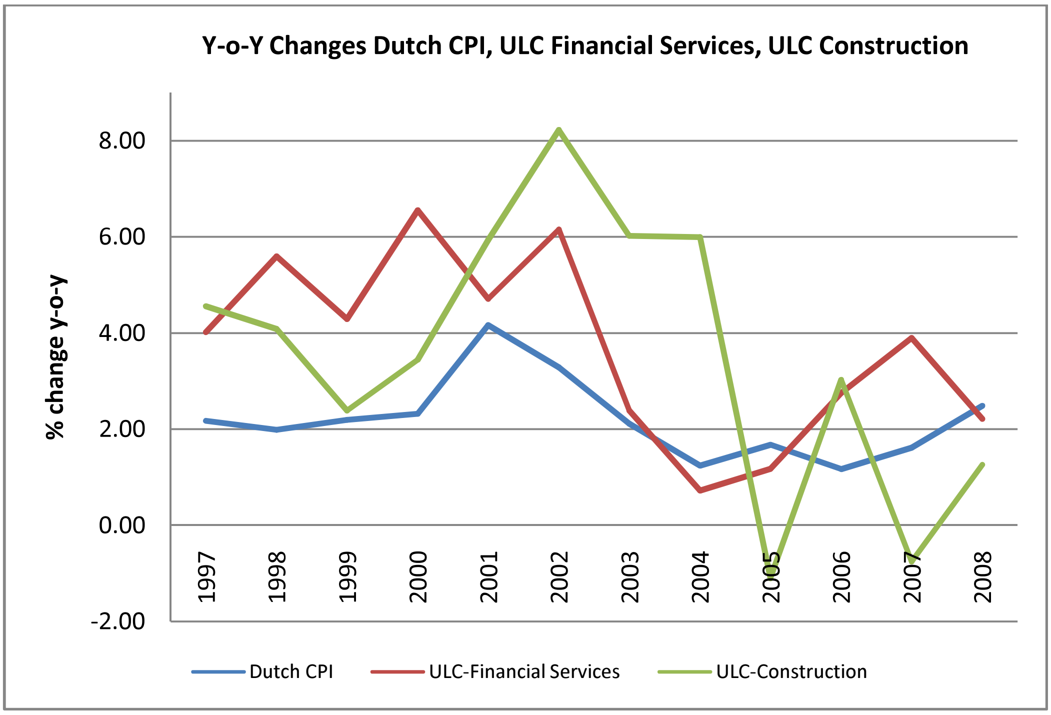
Wage inflation and pension fund indexation:

The Netherlands has one of the highest union coverage ratios in the world. The setting of yearly wages in the Dutch labour markets is a bargaining process between the union(s) and the employers (collectively known as the social partners). This bargaining process is the link between wage inflation and price inflation. The CBS website quotes the following: "The CPI reflects developments in the prices of goods and services which a consumers buys. It is an important measure for inflation and is used on a wide scale by the government and in trade and industry, *among other things in wage negotiations*, for *the indexation of* rents and *annuities* and to adjust tax tables".[5] There may be other factors also depending upon whether the negotiations are at the firm level or at the sector level. However it is reasonable to start with the assumption that wage inflation and price inflation in the Netherlands should have long/short associations.

Pension funds in the Netherlands index liabilities to either wage inflation or price inflation. Furthermore indexation is conditional on the health of the fund[6]. Since pension funds in the Netherlands can also be sector specific, like construction, civil servants, health care etc., the wage inflation rate used in such cases for indexation is sector specific. To get an intuitive feel of how wages evolve vis-à-vis price two charts are presented. One is for unit labour cost[7] (ULC) for financial services and the construction sector and the other for ULC of the total economy. The time varying correlation in Graph 6 captures the long-run relationship between wage measure using ULC total economy and prices using Dutch-CPI and shows that correlation tends to one over time. This implies that in the short-run, wage inflation and price inflation may diverge, but in the long-run they are strongly correlated. This is a very crude way to quantify the long-run relationship. We leave the details for later chapters.

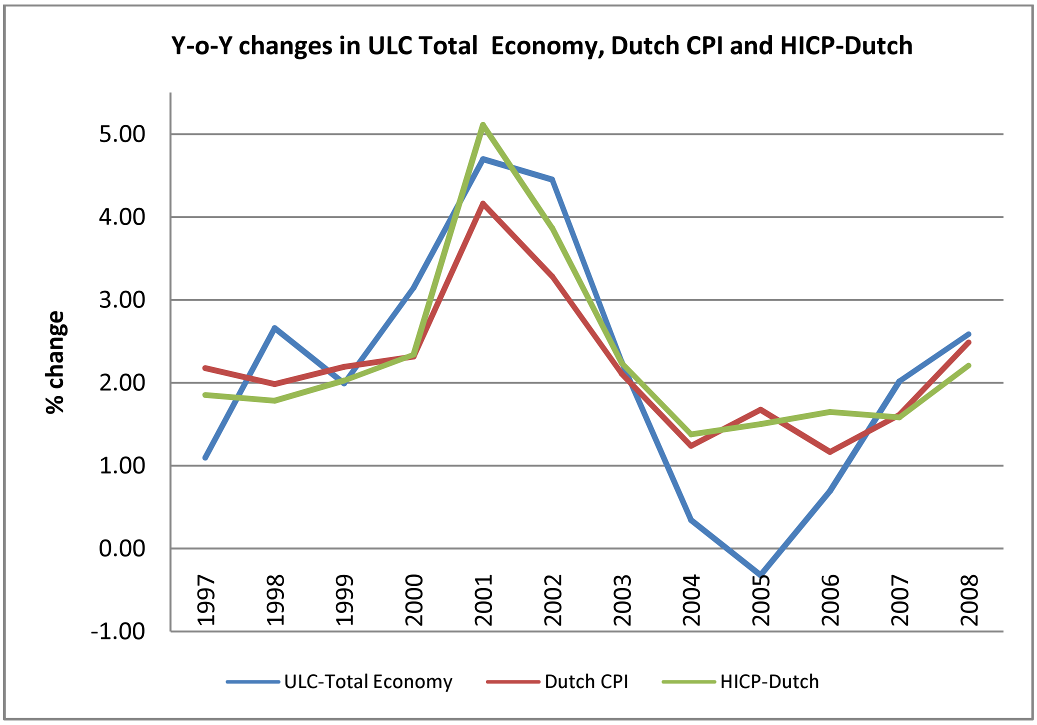
[4] Garcia and Rixtel (2007); Inflation Linked Bonds from central bank perspective, ECB Occasional paper series, No 62, June 2007.
[5] Emphasis added.
[6] Since indexation is conditional it is commonly referred to as indexation ambition else it would be a guarantee.
[7] For an exact definition of this indicator of wage see Chapter VI.

Graph 4



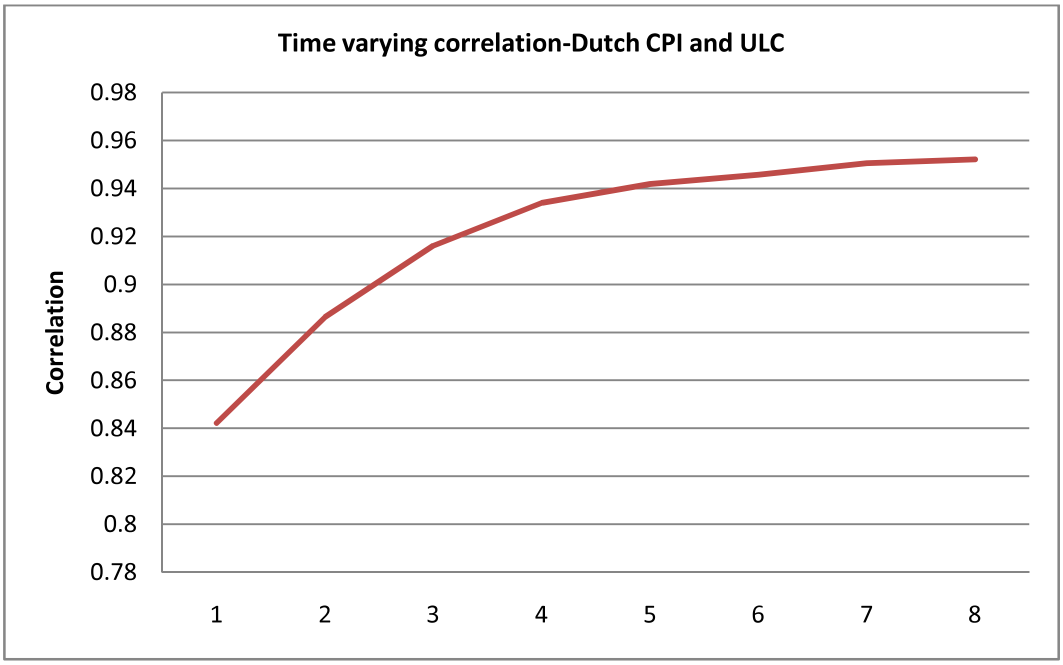
Source: OECDStats, Eurostats. (Base 2005=100)

Graph 5



Source: OECDStats, Eurostats. (Base 2005=100)

Graph 6



Source: OECDStats (Base 2005=100). The rolling correlations between changes in Dutch-CPI and ULC are obtained using an expanding window starting from 1 year, 2 year till 8 years

III

Economic of wage-price dynamics[8]

Traditional macroeconomic theory looks at wages and prices from a very different perspective. In macroeconomics, wages and prices are considered as demand side indicators and their interplay is a key determinant of the stabilization policy (monetary policy and fiscal policy). In case of pension funds the wage enters the liability side of the pension fund balance sheet because of the indexation ambition and price inflation on the asset side of the balance sheet if inflation linked derivatives are used (for hedging the ambition of wage indexation). The government stabilization policy targeting wages or prices affects the pension fund balance sheet directly. The dynamics of wages and prices at the macroeconomic level therefore have an important bearing at the pension fund level. **The purpose of this chapter is to summarise how traditional macroeconomic theory perceives the wage-price mechanism**. This sets the stage for the empirical survey on wage-price modelling in the next chapter.

Macroeconomic stabilization policies in the 1970s based on medium-term strategy have been a hallmark of industrialised countries. An important **assumption of this strategy is that there is no long run trade-off between price inflation and output**. This assumption is also known as 'the long run neutrality of money' or the 'natural rate hypothesis (NRH)'. An implication of this assumption is that an expansionary monetary policy may temporarily increase output and lower unemployment, but in the long run the economy will return to the natural rate of unemployment at a higher inflation. The success of the medium term strategy across developed countries has largely been mixed. While there is evidence for a NRH in the United States (US) there is no conclusive pattern in Europe. In an attempt to explain this phenomenon, it is natural to look at the wage-price mechanism. It is from this point in the macroeconomics literature that we start finding answers to the problem in hand.

The initial attempts to model wage-price dynamics started with the assumption that a long-run trade-off between inflation and output (Y) exists. The long-run trade-off can be described in a simple equation as under:[9]

$$\Delta P = a + b\Delta Y \; ; b > 0 \qquad \textbf{(3.1)}$$

The Old Keynesian model assumes the existence of equation (3.1) while the New Classical and subsequent models refute it.

The Old Keynesian Model:

Traditional Keynesian economics postulates the notion of slowly changing prices and wages. In the 'General Theory', Keynes (1935) assumed that the nominal wages are sticky which prevents the labour market from clearing but output prices are flexible. The model was static but with the discovery of the Phillips curve (1958) it was possible to make the model

[8] This chapter borrows from: Andersen (1989): Inflation and output: A review of wage-price mechanism, BIS Economic Papers No. 24-Jan 1989.
[9] All variables included in the equations are in natural logarithms. The purpose of the equations used henceforth is to bring the intuition of the theory. Hence we suppress the time subscript and avoid econometric issues here.

dynamic. The Phillips curve postulates a negative relationship between the changes in nominal wages (ΔW) and rate of unemployment (Δu), that is:

$$\Delta W = c - d\Delta u\,; d > 0 \quad \textbf{(3.2)}$$

By assuming that prices are set as a constant mark-up over trend unit labour cost (i.e. W-q where q is productivity)[10] and that there is a linear relationship between (excess) supply in labour market and product market [Okun, (1962)], it is possible to determine prices and wages simultaneously. Hence we have the two equations-the mark-up equation and the Okun equation:

$$P = eW - fq \quad ; e, f > 0 \quad \textbf{(3.3)}$$

$$\Delta Y = g - h\Delta u\,; h > 0 \quad \textbf{(3.4)}$$

Substituting (3.4) in (3.1), rearranging and substituting for Δu in (3.2) we can be get two simultaneous equations containing wage and price. Except for the US, the Okun equation (3.4) has been less stable for other countries and the price equation (3.3) has two shortcomings: First, in case of a small open economy the price changes will be affected by developments in exports and imports. Second, when prices are assumed as a constant mark-up over the trend unit labour cost, the real wages will rise with rising labour productivity. The distribution of factor income will therefore be constant, except for cyclical fluctuations. However outside the US, distributional issues have played a prominent role in modelling wage-price mechanism.

The old Keynesian model broke down in 1970. The aforementioned overlook of foreign prices was a major shortcoming of the model. The more serious criticism came from Friedman (1968) and Phelps (1968) who questioned the validity of the long-run permanent trade-off between price inflation and output i.e. equation (3.1). They suggested the absence of a long-run trade-off. Furthermore, they proposed that the Philips curve should be specified as a relationship between expected real wage changes and changes in unemployment. This meant that equation (3.2) is now specified as under, where P^e is the expected price level and depends on past price level that is expectations are formed adaptively.

$$\Delta W = i - j\Delta u + k\Delta P^e \quad \textbf{(3.5)}$$

The New Classical Model:

However the Phelps-Friedman assumption that prices are formed adaptively was criticised by Lucas (1972). Lucas argued that rational agents form expectations based on all available information. Secondly, the New Classical model assumes that (i) prices clear product market instantaneously and (ii) each firm is a price taker. Another feature of the New Classical model is that wage and price equations are influenced by changes in policy regimes. This sensitivity to policy regimes is also referred to as the Lucas critique [Lucas, (1976)]. Together with the above two assumptions and the Lucas critique, the model implies that anticipated changes in policy have no material impact on output. **The New Classical contribution therefore has been the idea of incorporating rational expectation and the insight that short run deviations can only occur if government policy is unexpected.**

[10] See Box 1 for some comments on this assumption.

New Keynesian Models:

The new Keynesians models of wage-price mechanism have retained the assumption of sticky prices and wages and non-clearing markets but have adopted the notion of **no long-run trade-off between inflation and output** as well as the New Classical assumption of rational expectations and optimising agents. The model is therefore characterised by the equations (3.3), (3.5) (appropriately modified) along with the following equation for rational expectations:

$$\Delta P^e = E(\Delta P / Given\ all\ information) \qquad \textbf{(3.6)}$$

Wage-price mechanism in Europe:

The NRH has been adopted in all models for wage-price mechanism. However as mentioned before the NRH was not been validated everywhere outside the US. Stabilisation policies in Europe were able to bring the inflation down but the unemployment remained above its initial level. The rate of inflation did not continue to decelerate as NRH would predict. Hence it is compelling to **assume that Europe still fits in the Old Keynesian model with a permanent long-run trade-off between price inflation and output.** The anomalous behaviour of unemployment led to many explanations for the European problem. The explanations rested on many kinds of hypothesis, to name a few- structural rigidity, hysteresis, real wage rigidities, output-price constraints and high population and labour force growth. None of the hypothesis completely explains the European problem because the European problem is a complex interplay of many factors.

Conclusion:

From the analysis we infer that the Old Keynesian model was a framework for understanding the long-run relationship. Equations (3.1) to (3.4) are all long-run equations to ensure consistency. But as soon as we drop the assumption of a long-run trade-off between output and prices, we fall in the domain of short-run analysis. Hence both **the Classical and the New Keynesian theories are short-run theories for wage-price dynamics** and all equations are short-run equations. However if equations (3.3) and (3.5) explain short-run behaviour then there is a reason to believe that in the long-run wage and prices are correlated. This approach has been adopted in the empirical literature to understand the long-run behaviour of wages and prices. This point is highlighted at the beginning of the next chapter.

IV

Empirical Literature on wage-price dynamics

In this chapter we briefly explain how the macroeconomic theory explained in the previous chapter was translated into empirics. The theory and its assumptions highlighted in previous chapter form the basis of these studies but readers will notice that application is not straight forward [See Box 1.] Another aim of this chapter is to understand how the price/wage equation has been estimated. The early work on modelling the wage-price relationship relied on time series analysis. The aim of these studies was to devise and understand stabilization policies in the 1980. It should be borne in mind that econometric theory of long-run analysis using co-integration did not exist at this time. Led by Gordon (1988), separate wage and price equations[11] were estimated of the following type:

$$\Delta p_t = h_0 + h_1\Delta(w_t - q_t) + h_2 G_t + h_3 S_{pt} + \varepsilon_{pt} \qquad \textbf{(4.1)}$$

$$\Delta(w_t - q_t) = k_0 + k_1\Delta p_t + k_2 G_t + k_3 S_{wt} + \varepsilon_{st} \qquad \textbf{(4.2)}$$

where all variables are in natural logarithms and where p_t is the price level, w_t the wage rate, G_t is the output gap i.e., the difference between nominal output and potential output, q_t is labour productivity, S_{Pt} is a supply shock on the price equation typically oil prices and S_{Wt} is a supply shock on the wage equation.[12] Equation (4.1) describes the price-mark-up behaviour. That is, price is set as a mark-up over the productivity adjusted wage and is affected by cyclical demand and supply shock variables. This equation therefore implies that, given other variables, productivity adjusted wage determines the price level. Equation (4.2), a Phillips curve, is the wage equation and states that, given other variables changes in wages are determined by changes in prices. **The wage-price behaviour described by the system of equation (4.1) and (4.2) implies that the long-run movements in wages and prices must be related**.

The above formulation is an application of the (expectations augmented) Phillips curve and the New Keynesian theory of wages and prices. Following this approach, Gordon (1988) showed for the US data that if price changes precisely mimic wage changes, then the wage equation is redundant. The wage changes do not contribute statistically to the explanation of price inflation, i.e. h_1=0, and price inflation depends on past inflation but not on past wage changes. The results by Gordon were in sharp contrast to earlier results which had found wages and prices to be related [See Barth and Bennett, (1975); Mehra, (1977).]

This contrasting finding was revisited by Mehra (1991). The author pointed to the fact that earlier studies had assumed that wages and prices contained deterministic trend only. Nelson and Plosser (1987) had questioned this assumption for macroeconomic time series. According to Nelson and Plosser, trend components of macroeconomic series also contain stochastic elements. Mehra (1991) confirms that variables measuring the rate of growth of wages and prices do not contain deterministic trends only but shared instead a common

[11] The system of equations (4.1) and (4.2) is adapted from Mehra (1991). Gordon's equation included lags of dependent variables on the RHS. Mehra supplemented this system of equation (4.1) and (4.2) with third equation $\Delta p_t^e = \sum_{j=1}^{n} \theta_j \Delta p_{t-j}$ with the Δp_t in RHS of (4.2) now replaced by Δp_t^e.

[12] The term supply shock refers to a sudden rise in the price of key industrial input. Prominently used in 1970s it refers to the rise in oil prices which resulted in a sudden rise in input cost and final consumer prices.

stochastic trend, that is, these variables are co-integrated as discussed in Engle and Granger (1987). A study similar to Mehra's (1991) on Irish data was performed by Fountas, Lally and Wu (1999). The study also reports presence of co-integration between wages and prices for Ireland. The presence of co-integration should be not ignored in equations (4.1) and (4.2) because a misspecification of the trend component can lead to an incorrect test of one's hypothesis [see Stock and Watson (1989)] and ignoring the error correction term leads to an omitted variables bias.

This two equation specification of the type in (4.1) and (4.2) was eventually replaced by a reduced form equation for prices or a single equation for the wage. The expectations augmented Phillips curve relationship relates wage growth to (expected) growth in prices, the unemployment rate and changes in unemployment. But we observe that in Equations (4.1) and (4.2) the unemployment variable does not enter as an explanatory variable. Besides many typical wage bargaining factors, which are highly relevant to the European labour markets, are also missing. It is because of this reason that an extended version of the traditional Philips curve equation that includes other explanatory variables like rate of income tax, social premium and labour productivity is popular in Europe [Graafland, (1991).] Theoretical justification of a 'bargaining augmented Philips curve' was given by Knoester and Van der Windt (1987).[13]The typical wage equation [see Broersma and Butter (1999)] has the following form:

$$\Delta w_t = a_0 + a_1 \Delta p_{c,t} + a_2 \Delta p_{y,t} + a_3 \Delta q_t + a_4 \Delta t_{p,t} + a_5 X + \varepsilon_t \qquad \textbf{(4.3)}$$

where all variables are in natural logarithms and where p_c is the consumer price level, p_y is the producer price level, w_t the wage rate, q_t is labour productivity, t_P are employer and employee taxes and social premiums; and X is a matrix containing other labour market variables like the unemployment rate.

A reduced form for the price equation, proposed by Gordon (1997), popularly known as the Triangle model for price inflation, has the following form:

$$\pi_t - \pi_{t-1} = a_1(u_t - u^*) + a_2 S_t + \varepsilon_t\,; \quad a_1 < 0,\, a_2 > 0 \qquad \textbf{(4.4)}$$

where π_t is the rate of inflation, u_t is the unemployment rate, u^* the natural rate of unemployment and S_t is the supply shock variable. In a small open economy setting with centralised bargaining framework, Petursson (2002) studied the wage-price formation for Iceland. Resorting to statistical techniques for non-stationary time series, their error correction model can be interpreted as a generalised version of the Philips curve for wages and prices, derived from the New-Keynesian theories of sticky wages and prices. This model is also a generalised version of the Gordon's (1997) triangle model for inflation.

Studies cited till now have used time series analysis to test the underlying theory. The availability of long macro time series for several countries has prompted some macroeconomists to test the theory using panel data. **This approach analyses the existence of a common average Phillips curve for a set of countries**. DiNardo and Moor (1999) use panel analysis to examine nine OECD countries. They used Ordinary Least Square (OLS) and Generalised Least Square (GLS) to conclude that there is "a remarkably

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[13] The economic literature has another approach which builds on the work of Sargan (1964) where the wage equation is derived from a microeconomic theory of wage bargaining.

robust relationship between relative inflation and relative unemployment". Turner and Seghezza (1999) also used a panel analysis on twenty one OECD countries over a period of from early 1970 to 1997. They used the technique of Seemingly Unrelated Estimation (SUR) to analyse the panel data and conclude that there is "strong support [for a] common Philips curve among 20 member countries of the OECD". More recently, Bjornstad and Nymoen (2008) tested the hybrid New Keynesian specification on a panel of twenty OECD countries. Their aim however was not to test a particular strand of the economic theory. In the coming pages we turn our attention to the econometric theory for panel unit root and panel co-integration testing, estimation methodology and interpretations of results. [See Box 2].

Box 1

Critical remarks on wage and price equations

The wage-price mechanism is modelled as a system of simultaneous equations. Thus both the wage equation and the price equation are equally important. However the theoritical literature has concentrated more on the wage equation [see Sargan (1964), Knoester and Van der Windt (1987), Graafland (1992), Manning (1993)] than on the price equation. Most of the studies conclude that prices are fully reflected in wages, that is, the mark-up assumption is reasonable, albeit with certain lag. Therefore as regard to price equations many features are disturbing and seriously hamper the study of the wage-price mechanism. Nordhaus, (1972) in a survey paper mentions the following major shortcomings with regard to price equations:

1) Most specifications and interpretations have proceeded without the benefit of a formal theory. As a result, the implicit elasticities of price with respect to different costs are difficult to understand.

2) Very little is know about the structure of the impact of demand on prices apart from the effect through the unemployment rate on wage.

3) Authors have been very casual about whether they measure prices in level or in first difference.

4) On econometrics; very little attention has been given to structure of errors.

5) Most authors estimate prices as a mark-up on unit labour cost and very few include other costs such as cost of capital, raw material etc.

Box 2

Some insights and approaches to the problem

In light of the above discussion following are some of the main considerations:

1. The new Keynesian theory postulates that there should be a long-run relationship between wage and price. The empirical literature provides some evidence that a long run relationship exists and hence for the reasons provided in the introduction of this thesis, pension fund must not ignore this fact.

2. The strand of theory which predicts this relationship is not that relevant, so long as the long term relationship is stable. This implies the causality that price determines wages or vice versa is not that important. [This argument is valid for pension funds but not true in general. For instance causality is important in macroeconomic policy making.]

3. Quantification of the long run relation rests on availability of long time series. Since HICP indices are available from 1996 the time series approach using HICP is untenable. Hence an indirect approach is required. [See point 6 below.]

4. Mixed results on wage-price relations in a European context require that time series techniques be approached with caution. Studies point that power of tests to distinguish unit roots and co-integration are limited with time series. Panel analysis therefore is the logical choice.

5. The panel data should include countries similar to the Netherlands. By this we mean we cannot include the US in our panel because labour markets in the US differ significantly (e.g. labour market rigidity, union coverage etc.) from labour markets in continental European countries such as the Netherlands.

6. To summarise; the problem in hand can be reformulated as – if an average or common relationship between wages and conventional price indicators like CPI exists, then the emergence of HICP should not change the very nature and dynamics of this relationship. The average relationship can then be used to understand the wage-price mechanism for individual countries. This should be the case at least in EU countries because HICP is standardized across member countries and there is a unified market on HICP derivatives.

V

Econometric Methodology

Co-integration analysis for panel data has three steps.[14] The first step is to ascertain whether the variables under study are non-stationary, i.e., whether the time series have a unit root. If the series under study are non-stationary, the second step requires establishing the presence of a long-run relationship, statistically: the co-integration vectors. The third step is the estimation of the long-run model and the resultant Vector Error Correction model (VECM) for the short-run dynamics. In the following paragraphs we describe the hypothesis testing procedure used in the first two steps and the estimation technique for the third step.

Panel Unit root tests:

The Dickey-Fuller test (DF) and the Augmented Dickey-Fuller (ADF) test are two most prominent tests used in the time series literature to test for non-stationarity. The ADF test involves estimating the equation:

$$y_t = \alpha + \beta t + \gamma y_{t\text{-}1} + \sum_{j=1}^{k} \theta_j \Delta y_{t\text{-}j} + \varepsilon_t \qquad \textbf{(5.1)}$$

The series is said to have unit root if γ=1. The limiting distribution of this test is not standard and simulations are used to determine the critical values. Many problems (size distortion and low power) have been reported for DF and ADF tests.

Low power of unit root tests in time series has often been cited as the reason for the use of panel data that improves the power by using both time dimension and cross section. Levin, Lin and Chu (LLC) extend the ADF equation (5.1) to panel data to test the hypothesis that each individual time series contains a unit root against the alternative that each time series is stationary. LLC estimate the equation-

$$\Delta y_{it} = \alpha + \beta t + \rho y_{it\text{-}1} + \sum_{L=1}^{k} \theta_{iL} \Delta y_{it\text{-}L} + \varepsilon_{it} \qquad \textbf{(5.2)}$$

where $\rho = \gamma$-1, i = 1,2 ... N, and the inclusion of a constant and a trend or both depends on a case to case basis. Since the lag order k is unknown, LLC suggest a three step procedure to determine k. They proposed a t-statistic t_ρ^* which is asymptotically distributed N(0,1). LLC assume independence across cross-section units in order to apply the law of large numbers when aggregating over N. O'Connell (1998) showed that LLC suffered from significant size distortions in the presence of correlation among cotemporaneous cross-sectional error terms.

LLC test is restrictive in a way because it requires ρ to be homogeneous across all i. Im, Pesaran, and Shin (1997) (IPS) allows for heterogeneous co-efficient for $y_{i\,t\text{-}1}$ i.e. replacing ρ by ρ_i in Equation (5.2). IPS proposed an average of ADF tests where the error is serially correlated. The null hypothesis is H_0: ρ_i=0 for all i against the alternative hypothesis that some individual series have a unit root, i.e., H_1: ρ_i<0 for i=1, 2, ..., N_1 and ρ_i=0 for i=N_1+1, N_1+2, ..., N. IPS t-statistic is an average of the individual ADF t-statistics. IPS test also assumes independence across cross-section units in order to derive the asymptotic

[14] Please refer to Appendix A for a formal definition of co-integration in time series as a prelude to this chapter.

properties of their statistics. IPS showed that the asymptotic distribution of the statistics tends to N(0,1) when T tends to infinity and N tends to infinity (the sequential limits.)

Maddala and Wu (1999) and Choi (1999a) proposed a non-parametric Fisher test by combining the p-values from individual unit root tests. The test statistic is

$$P = -2\sum_{1=1}^{N} ln\,(p_i) \qquad \sim \chi^2\,(2N) \tag{5.3}$$

This combines the p-vales from the unit roots from each cross section i to test the unit root in panel data. -2ln (p_i) has a Chi-squared distribution with 2 degrees of freedom. Thus P is distributed as Chi-squared with 2N degrees of freedom when T_i is large relative to N. Maddala and Wu argue that Fisher's test, like the IPS test, relaxes the restrictive assumption of LLC test. However, **Fisher's test has an advantage over IPS test in that it does not require a balanced panel**. Also, Fisher's test allows for variable lag lengths across individual ADF regressions. Because the IPS test and Fisher test combine the significance of different independent tests, they are directly comparable. Furthermore "Fisher's test is an exact test while IPS test is asymptotic. But the asymptotic validity of the tests depends on different conditions. For the IPS test the asymptotic result depends upon N going to infinity while for Fisher test they depend on T going to infinity" [Maddala and Wu, (1999)]. Choi (2001) studied the small sample properties of IPS test and Fisher test. He concludes that the empirical size of the IPS test and Fisher test are reasonably close to their normal size of 5% when N is small. In terms of the size-adjusted power, the Fisher test seems superior to the IPS test.

There are other unit root tests in the literature which we do not use for this study. Hadri (2000) proposed a residual based Lagrange Multiplier test. The empirical size of this test is close to the 5% level for sufficiently large N and T. In our case N is not very large. Pesaran (2004) suggested a simple test for unit root that account for cross-sectional dependence for errors in panel with short T and large N. This is typical for a micro panel but not for a macro panel where T is large relative to N.

Panel Co-integration test(s):

Panel co-integration tests can be constructed using two approaches. The first approach is to base the tests on Engle and Granger (1987) methodology whereby residuals from the least square estimate of co-integration regression are subjected to unit root test. Notable contributions using this approach are Kao (1999), McCoskey and Kao (1998), Pedroni (2000, 2004). The second approach proposed by Westerlund (2007) is a structural based test for co-integration rather than based on residual dynamics. This test is designed to test for co-integration by testing the significance of the error correction term in the (conditional) error correction model. The following paragraphs describe some of these tests. The most relevant for our purpose are covered in detail.

<u>Residual based co-integration tests</u>: Like panel unit root tests, the tests for co-integration have also been inspired by the search for more powerful test. Hence the early tests for co-integration are an extension of time series. Consider the simple fixed effect model:

$$y_{it} = \alpha_{\mathrm{i}} + \beta x_{\mathrm{it}} + \gamma t + \varepsilon_{it}, \tag{5.4}$$

where y_{it} and x_{it} are I(1) and under the null hypothesis of no co-integration. For only a constant term and no trend, i.e., γ=0 in (5.4), Kao (1999) proposed DF and ADF type unit root tests on residuals, ε_{it} to test the null hypothesis of no co-integration. This entails running unit root tests on the residual from the fixed effect model and testing the H_0: ρ=0. Kao proposed four statistics based on DF test and one statistics using the ADF test. The ADF statistic uses the ADF equation of the type shown above on fixed effect residuals. All five statistics converge to standard normal distribution on application of sequential limit theory. McCoskey and Kao (1998) derived a residual based Lagrange Multiplier test *for the null hypothesis of co-integration*. For this test (with null of co-integration) it is essential to use an asymptotically efficient estimation technique to estimate the long-run equation (5.4) and obtain the residuals. Suggested efficient techniques are the fully modified ordinary least square (FMOLS) of Phillips and Hensen (1990) and the dynamic least square (DOLS) proposed by Saikkonen (1991) and Stock and Watson (1993). We explain DOLS method in some detail later in this chapter. Pedroni (1997a, 2000, 2004) also proposed seven residual based test statistics for the null of no co-integration. Pedroni's co-integration regression equation is -

$$y_{it} = \alpha_i + \beta_i x_{it} + \gamma_i t + \varepsilon_{it} \qquad \textbf{(5.5)}$$

Notice that the slope co-efficient β_i is allowed to vary across cross-section units making (5.5) the most general specification among the residual based tests for co-integration. The residuals so obtained from equation (5.5) are tested for unit root. Of the seven statistics proposed, four are panel statistics, i.e., they test the hypothesis for co-integration for the panel as a whole. The hypothesis tested is: H_0: $\rho_i = 0$ for all i against H_1: $\rho_i = \rho < 1$ for all i. The remaining three statistics are group statistics which test for co-integration across cross-sections. Hence, rejection of the null of no co-integration using a group statistic implies evidence of co-integration for at least one cross-section. **Since we are interested in average relationship across countries, panel statistics are more relevant for our case**. Furthermore, of the four panel statistics we report the Panel-ADF statistics only to maintain continuity from our previous discussion. The Panel-ADF statistic uses the ADF equation to test for the unit root on the residuals from Equation (5.5). The intuition for this test is exactly the same as described for the other residual based test above [For more details readers are invited to see Pedroni (1997a).]

Structural co-integration tests: The residual based test requires that the long-run co-integrating vector for the variables in levels is equal to the shot-run adjustment process for the variables in the first differences. Kremers *et al.* (1992) refer to this as the common factor restriction and show that its failure can result in significant loss of power for residual based tests. Taking this limitation of residual based test as the starting point, Westerlund (2007) proposes a structural based test for co-integration rather than one based on residual dynamics and therefore does not impose the any common factor restrictions. The underlying idea is to test for the absence of co-integration by determining whether there exists an error correction for individual panel members or for the panel as a whole. Consider the following error correction model, where all variables in levels are assumed to be I(1):

$$\Delta y_{it} = \delta_1 + \delta_2 t + \alpha_i\left(y_{it-1} - \beta_i\, x_{it-1}\right) + \sum_{j=1}^{p_i} \vartheta_{ij} \Delta y_{it-j} + \gamma_{i0} \Delta x_{it} + \sum_{j=1}^{p_i} \gamma_{ij} \Delta x_{it-j} + e_{it} \qquad \textbf{... (5.6)}$$

The test requires estimating the parameter α_i, which can be interpreted as the speed of adjustment to the long-run equilibrium. Equation (5.6) is estimated using OLS. The test is designed to test for co-integration by testing the significance of the adjustment term in Equation (5.6). **The null hypothesis for this test is no error correction, i.e., α_i=0 and its rejection implies rejection of the null of no co-integration.**

The intuition for the test is as follows: the error correction model will be stable only if the term $\boldsymbol{y}_{\text{it-1}} - \boldsymbol{\beta}_{\text{i}}\ \boldsymbol{x}_{\text{it-1}}$ is stationary. Any deviation from the long-run leads to correction by a factor $-2 < \alpha_i < 0$. If $\alpha_i<0$, then there is error correction, which implies y_{it} and x_{it} are co-integrated. Where as if $\alpha_i=0$ error correction term is irrelevant and there is no co-integration. As in Pedroni's test, there are two sets of statistics- two group statistics and two panel statistics. We concentrate on panel statistics only. The panel statistics P_α and P_τ for i^{th} cross-section are defined as:

$$P_\tau = \frac{\hat{\alpha}}{SE(\widehat{\alpha})} \qquad\qquad P_\alpha = T\hat{\alpha}$$

They test the hypothesis: H_0: $\alpha_i = 0$ against H_1: $\alpha_i = \alpha < 0$ for at least some i, suggesting that a rejection should be taken as evidence of co-integration for the panel as whole. All the four statistics converge to standard normal distribution[15] on application of sequential limit theory and are consistent. Westerlund tests are fairly general because they are valid for weekly exogenous x_{it}, varying lags across cross-section permitting heterogeneous serial correlation structure, cross-sectional dependence by using the bootstrap approach which makes inferences possible under fairly general form of cross-sectional dependences.[16]

Estimation and inferences in static and dynamic panel models:

The method for estimating the coefficients in the long run equation are different from those for estimating the coefficients of the short-run model. This difference arises because short-run model is dynamic in nature that is it involves a lag of the dependent variable. The purpose of this section is to describe in brief, the methods for estimating the long-run equation and the short-run equations.

Long-run equation: For panel co-integrated regression equations, the asymptotic properties of the estimators and their associated statistical tests are different from those in time series. Chen, McCoskey and Koa (1999) have investigated the finite sample properties of OLS estimators, t-statistics, the bias corrected OLS estimator and bias corrected t-statistics. They conclude that bias-corrected OLS estimator do not improve over the OLS estimator. Chen et al suggest using alternatives such as the fully modified OLS (FMOLS) and the dynamic OLS (DOLS) estimators to estimate the long-run co-integrating vector(s). Kao and Chiang (2000) also investigated OLS, FMLS and DOLS for co-integrated panel regressions and find that – (i) the OLS estimator has a non negligible bias in finite sample, (ii) the FM estimator does not improve over the OLS estimator in general and (iii) the DOLS estimator may be more useful

[15] The normality for P_α at first sight looks counterintutive. But Banerjee *et al.* (1998) proved that this form of statistics also has a limiting normal distribution under the alternate hypothesis.

[16] Westerlund's tests have been executed using the xtwest command for Stata made available by the author. xtwest is not an official Stata command. Please refer Persyn, D. and J. Westerlund. 2008. Error Correction Based co-integration tests for Panel Data. Stata Journal 8 (2), 232-241. Results are reported in the following chapter.

than OLS or FM estimator in estimating the co-integrated panel regression. Hence we use DOLS for this study and describe this method in some detail below.

Mark and Sul (2003) extended the time series DOLS method of Saikkonen (1991) and Stock and Watson (1993) to the panel data. The DOLS for the simple fixed effect is explained below [For other cases the readers are invited to see the paper by Mark and Sul (2003).] Consider the long-run model:

$$y_{it} = \alpha_{\mathrm{i}} + \beta x_{\mathrm{it}} + u_{it} \quad \textbf{(5.7)}$$

$$x_{it} = x_{it\text{-}1} + \varepsilon_{it} \quad \textbf{(5.8)}$$

Assume that u_{it} is correlated with at most p_i leads and lags of ε_{it}. To control for endogeneity DOLS entails regressing u_{it} on the p_i leads and lags of $\varepsilon_{it.}$ Let δ' be the vector of regression coefficients from this regression and $z_{\mathrm{it}} = (\Delta x'_{it-p_{\mathrm{i}}} \dots, \Delta x'_{it} \dots\dots, \Delta x'_{it+p_{\mathrm{i}}})'$ denote the vector of all the lags and leads i.e. we have the regression equation, $u_{\mathrm{it}} = \delta_i' z_{\mathrm{it}} + u_{it}^*$. Substituting for u_{it} in equation (5.8) we have:

$$y_{it} = \alpha_{\mathrm{i}} + \beta x_{\mathrm{it}} + \delta_i' z_{\mathrm{it}} + u_{it}^* \quad \textbf{(5.9)}$$

Taking deviations from mean for the respective time series in equation (5.9), the DOLS estimate of β denoted by β_{NT} is given by:

$$\beta_{\mathrm{NT}} - \beta = [\sum_{i-1}^{N} \sum_{t=1}^{T} \tilde{q}_{it}\, \tilde{q}_{it}']^{-1} \{\sum_{i=1}^{N} \sum_{t=1}^{T} \tilde{q}_{it}\ \ \tilde{u}_{it}^*\}$$

Where 'tilde' denotes that observation is deviation from mean and the vector q is appropriately defined and contains mean deviations of x_{it} and the p_i lags and leads and zeros[17]. When T tends to infinity and N tends to infinity the DOLS estimator is normally distributed. The appropriate lag length is selected using AIC or BIC criterion.

Short-run model: The VECM model entails estimating a fixed effect model that contains lags of the dependent variables. This makes the model dynamic. The following short note gives a brief overview of the methods used for this type of analysis. We keep the discussion concise because the estimation techniques for the dynamic panel models have developed for panels with N large relative to T. However we use a panel where T is large relative to N thus rendering a detailed discussion inappropriate. The literature for dynamic panels with T large relative to N is still developing and the current theory on the subject is a set of recommendations.

A simple dynamic fixed-effect model is specified as under:

$$y_{it} = \alpha_i + \beta x_{it} + \gamma y_{it\text{-}1} + \varepsilon_{it}; \quad |\gamma| < 1 \quad \textbf{(5.10)}$$

The basic problem with Equation (5.10) is that y_{it} is a function of α_i and it immediately follows that y_{it-1} is also a function of α_i. Therefore y_{it-1} is correlated with the error term. Thus the OLS estimator is biased and inconsistent. One way to get rid of the fixed effect is to estimate the model in first differences. This method was proposed by Anderson and Hsiao (AH) (1981).

[17] To put it simply, we estimate Equation (5.9) using OLS and choose that lag and lead length that minimises BIC.

However differencing only removes the fixed effect. The differenced right side variable is still correlated with the differenced error term. The problem at hand is how to obtain an appropriate estimate of the coefficients for panel models as described in Equation (5.10)?

Nickel (1981) showed that using a least square dummy variable (LSDV)[18] to estimate the fixed effects models like (5.10) results in biased estimates when time dimension of panel T is small. Nickel (1981) derived an expression for the bias of γ and shows that the bias approaches zero as T tends to infinity. For macro panel where T is large relative to N like ours, there are two important considerations:

(i) How large should T be before the Nickel bias can be ignored?
(ii) Most of the estimation techniques available for dynamic panel models have been derived for large N and small T. For macro panel like ours where T is large relative to N, different estimation techniques can produce very different results. Thus the choice of estimator in our case is crucial as competing estimators have generated different results [See Judson and Owen (1996, 1999).]

Several estimators have been proposed to estimate equations of the type (5.10) when N is large relative to T. These methods are listed as follows:

(i) The **Instrumental variable estimator of Anderson and Hsiao** (1981) (AH) on transformed model in first differences to eliminate unobserved fixed effects and then use the second lags of dependent in differences (or in levels) as instrument for one-time lagged dependent variable. AH estimation method leads to consistent but not necessary efficient estimate of the parameters in the model because instrumental variable does not make use of all available moment conditions [Ahn and Schmidt , (1995).]
(ii) One step **generalised method of moments (GMM) estimator of Arellano and Bond** (AB) (1991)(GMM1) and two step GMM estimator of AB (1991)(GMM2). This method is generalization of AH and is more efficient than AH because it uses all the moment conditions. The only difference between AH and AB is that AB proposes using lags of dependent in levels as instrument for one-time lagged dependent variable[19].
(iii) A **bias corrected LSDV** (LSDVC) derived by Kiviet (1995).

We do not explain the details of all these estimators here as it involves matrix equations. Interested readers are invited to refer to the paper by Judson and Owen (1996, 1999) on critical review of estimation techniques for dynamic macro panel model and their relative merits and demerits. We summarise findings/recommendations by Judson and Owen on estimation of dynamic macro panel below:

[18] LSDV is also known as the within estimator. Within estimation entails first subtracting the time averages for each variable and then estimating the parameters β and γ by OLS. Subtracting the time averages, results in fixed effect α_i dropping out of Equation (5.10).
[19] The logic of obtaining the instruments in AB GMM is as follows: Consider the model (5.10) with no x_{it} term. Taking first difference we have $y_{it} - y_{it-1} = \gamma(y_{it-1} - y_{it-2}) + (\varepsilon_{it} - \varepsilon_{it-1})$. Observe that for t=3, y_{i1} is valid instrument since it is highly correlated with (y_{i2}-y_{i1}) but not with (ε_{i3}-ε_{i2}). Similarly for t=4 y_{i1} and y_{i2} are valid instruments. Hence we proceed in this fashion to get all instruments. For equation with x_{it} we proceed as follows: if x_{it} is strictly exogenous and all x_{it} in Equation (5.10) are correlated with α_i in Equation (5.10), then all x_{it} are valid instruments in for first difference of Equation (5.10) in levels or equivalently in first differences also [Also see Table 8 footnotes.]

(i) For macro panel data one should not dismiss the Nickel bias as insignificant, even when the time dimension is as large as 30. With T=30 the bias can be as much as 20% of the true value of the co-efficient of interest.
(ii) For panels of all sizes, a corrected LSDV generally has low RMSE. When the panel is unbalanced LSDVC may not be appropriate and an alternative may be needed.
(iii) The one-step GMM1 estimator outperforms two-step GMM2. Using a restricted GMM i.e. restricting the number of instruments to say 3 or 5 does not reduce the performance of this estimator.
(iv) After a detailed Monte Carlo analysis their recommendations are as under for dynamic macro panel models: LSDVC is the best choice but practical consideration may limit its applicability. GMM is second best solution. AH is a computationally simple option to execute. (See Table 2)

Table 2: Recommendation by Judson & Owen

	T<=10	T=20	T=30
Balanced Panel	LSDVC	LSDVC	LSDVC
Unbalanced Panel	GMM1	GMM1 or AH	LSDV

Judson and Owen study did not explicitly cover unbalanced panels in their analysis. The most recent contribution to estimate a dynamic (unbalanced) panel model is by Bruno (2005). Kiviet (1995) and Bun and Kiviet (2003) had proposed an asymptotic method to obtain an approximation for small sample bias of the LSDV estimator; Bruno (2005) extents their approach to the case of an unbalanced macro panel. Bruno (2005) specifies the model in equation (5.10). The method requires estimating the coefficients using the current theory (i.e. using LSDV, AH or GMM1) and then approximating the bias that arises due to unbalanced macro panel nature of the data. Bruno (2005) proposed three separate formulas for approximating the bias. This bias is then subtracted from the initial estimate obtained using LSDV, AH or GMM1. His Monte Carlo study uses two panel case (N,$\bar{T}$) = (20, 20) and (10, 40) where the second argument in brackets is average group size. The results of the Monte Carlo study are summarised below:

(i) The bias for both γ and β is decreasing in $\bar{T}$.The bias for γ is also decreasing in unbalancedness
(ii) While increasing $\bar{T}$ is always beneficial in reducing the LSDV bias, reducing unbalancedness at the expense of time observation, for given N and $\bar{T}$ may instead exacerbate the bias.[20]

[20] Bruno (2005) method has been executed using the xtlsdc command for Stata made availbe by the author. xtlsdvc is not an official Stata command. Please refer to Bruno, G.S.F. 2005b; ‘Estimation and inference in dynamic unbalanced panel data models with a small number of individuals’, CESPRI WP n.165 , Università Bocconi-CESPRI, Milan. Results are reported in the following chapter.

To summarise this section, the theory for estimating the short-run model for a macro panel data is still under development. Very little is known about the properties of the existing micro panel estimators when they are applied to macro panels and in particular to unbalance macro panels. The most recent work for the macro panel uses existing micro panel estimators or tried to correct the bias/error in micro panel estimator that may arise because the estimator is applied to macro panel. We will highlight in next chapter appropriately, when we present our results on how different authors have applied the above listed methods to estimate the panel VECM.

VI

Data Description, Methodology and Findings

Data definition and source:[21]

To analyse the wage-price formation in the Netherlands, we employ the panel co-integration and VECM technique on a panel of twelve EU countries that are included in the HICPxT EU index. The data has annual observations on wages and prices for the countries- Austria, Belgium, Finland, France, Germany, Ireland, Italy, the Netherlands, Portugal, Spain, Greece and Luxembourg. Some observations are absent for some of the countries and hence **the panel is an unbalanced panel dataset**. The following table gives an overview of the span of the data over time for each country:

Country	Frequency	Year
Austria	Yearly	1970-2008
Belgium	Yearly	1970-2008
Finland	Yearly	1970-2008
France	Yearly	1960-2008
Germany	Yearly	1970-2008
Greece	Yearly	1970-2008
Ireland	Yearly	1976-2008
Italy	Yearly	1970-2008
Luxembourg	Yearly	1970-2008
Portugal	Yearly	1970-2008
The Netherlands	Yearly	1969-2008
Spain	Yearly	1970-2008

Wages and prices are central variables for this study. Definitions for the proxies for each of these are described as follows:

Price (p): Consumer price index (2005=100) for all items from OECD data base. The index is not seasonally adjusted.

Wages (w): Unit labour cost (ULC) index (2005=100) from OECD data base. OECD metadata defines this indicator as – "unit labour costs measure the average cost of labour per unit of output. They are calculated as the ratio of total labour costs to real output, or equivalently, as the ratio of average labour costs per hour to labour productivity (output per hour)." The total labour cost is defined as – "annual total labour costs is compensation of employees (COE)… adjusted for the self employed by multiplying COE by the ratio of total hours worked by all persons in employment to total hours worked by all employees of businesses. This target variable covers a significant part of total labour costs such as wages and salaries; bonuses; payments in kind related to labour services (e.g. food, fuel, and housing); severance and termination pay and employers' contributions to pension schemes, casualty and life insurance and workers compensation."

[21] See Appendix B for country-by-country plot of the raw data.

Diagnostics: Panel unit root and co-integration:

As a first step, we perform the diagnostics for unit root and co-integration in wages and prices. We report the IPS test and the Fisher ADF test for the unit root as they are more relevant for our purpose. For the co-integration we report Pedroni's panel-ADF test and Westerlund panel tests because they test for the presence of co-integration for the panel as a whole. The panel unit root tests for wage and prices are reported in Table 3.

Conclusion unit root test: Two variations of Equation (5.2) were tried, one with a constant and second with a constant and a trend. **For the second variant both the ULC and CPI are non-stationary in levels i.e. we are unable to reject the hypothesis ρ_i=0 for all i= 1,2 ...12 but rejected the hypothesis ρ_i=0 for all i=1,2,...12 in first differences. Hence we conclude that both ULC and CPI are integrated of order one**. The model with only a constant is unable to establish that ULC and CPI are integrated of order one. Should we include a trend is the question of whether the theory dictates such an inclusion. For CPI the possible reasons could be that there is no deflation (negative inflation) observed in the sample period. Second central banks have tried to maintain price stability by keeping inflation stable at around a non-zero value say 2%. Hence we include a trend in both CPI and ULC to account for this data specific problem.

Table 3: Panel Unit Root Tests

Model	$\Delta y_{it} = \alpha + \rho_i y_{it-1} + \sum_{L=1}^{k} \theta_{iL} \Delta y_{it-L} + \varepsilon_{it}$	$\boldsymbol{\Delta y_{it} = \alpha + \beta t + \rho_i y_{it-1} + \sum_{L=1}^{k} \theta_{iL} \Delta y_{it-L} + \varepsilon_{it}}$	
	First difference	**Levels**	**First difference**
IPS Test			
Price (CPI)	0.045	0.403	0.002
Wage (ULC)	0.000	1.000	0.000
Fisher Test			
Price (CPI)	0.083	0.277	0.000
Wage (ULC)	0.000	0.854	0.000

Note: *Figures reported above are p-values. For both the tests H_0: ρi=0 for all i, H_1: <0 for i=1, 2 ...N1 and ρi=0 for i=N_1+1, N_1+2, ...N*

The presence of cross-sectional correlations among errors can affect the size of panel unit root test because asymptotic distributions are derived based on the independence assumption. To account this we had demean the both series.[22] We therefore conclude that there is evidence that the CPI series is I(1) and ULC series is I(1). Hence we can perform the co-integrations test to test the existence of a long run relationship.

Conclusion panel co-integration: We find that the wage and price series are co-integrated in line with the prediction of economic theory. Westerlund test concludes evidence for co-integration between wage and prices. **Both the Westerlund panel statistics reject the null of no co-integration at 1%, suggesting that there is strong evidence for a possible**

[22] Demeaning entails first subtracting the cross-sectional averages from the series. Hence for each time period we compute the mean of the series across panels and subtract this mean from the series. Levin, Lin, and Chu suggested this procedure to mitigate the impact of cross-sectional dependence. O'Connell (1998) reports gross oversizing in the context of testing for Purchasing Power Parity and also suggest cross-sectional demeaning.

common relationship between wages and prices for the 12 countries under study. Pedroni's panel-ADF statistics reject the null hypothesis at 10% significance.[23]

Table 4: Panel Co-integration Tests

Pedroni Test	
$\mathbf{y}_{it} = \boldsymbol{\alpha}_i + \boldsymbol{\beta}_i\mathbf{x}_{it} + \boldsymbol{\gamma}_i\mathbf{t} + \mathbf{e}_{it}$	
Panel ADF Satistics	0.080
Westerlund Test	
$\Delta\mathbf{y}_{it} = \boldsymbol{\delta}_1 + \boldsymbol{\delta}_2\mathbf{t} + \boldsymbol{\alpha}_i(\mathbf{y}_{it-1} - \boldsymbol{\beta}_i'\mathbf{x}_{it-1}) + \sum_{j=1}^{pi}\boldsymbol{\alpha}_{ij}\Delta\mathbf{y}_{it-j} + \sum_{j=0}^{pi}\boldsymbol{\gamma}_{ij}\Delta\mathbf{x}_{it-j} + \mathbf{e}_{it}$	
P_t	0.000
P_a	0.000

Note: *Figures reported above are p-values.*
Pedroni panel tests H0: $\rho_i = 0$ for all i against H1: $\rho_i = \rho < 1$ for all i
Westerlund panel tests H_0: $\alpha_i = 0$ against H_1: $\alpha_i = \alpha<0$ for at least some i

It is surprising that Pedroni's test give very week evidence for co-integration. This week evidence for co-integration using Pedroni's test is in line with the result reported by BjØrnstad and Ragnar (2008) for same OECD data for their estimate of the New Keynesian Phillips curve. Although the single most cited rational for using the panel tests is to increase the power of the test, many studies such as Ho (2002)[24] have also failed to reject the null hypothesis of on co-integration using residual based tests even when co-integration is strongly suggested by the theory. Westerlund (2007) provides the explanation that the residual based test requires that the long-run co-integrating vector for the variables in levels equal to the short-run adjustment process for the variables in differences. Kremers *et al.* (1992) refer to this as the common factor restriction[25] and show that its failure can result in significant loss of power for residual based tests [See Footnote 25.]

Long-run model:

The long run wage-price equilibrium relation is assumed to have the following linear form:

$$w_{it} = \alpha_i + \beta_i p_{it} + \varepsilon_{it} \qquad \textbf{(6.1)}$$

where i = 1,2, ..., 12 and t = 1, 2, ..., T_i and variables are included in natural logs. We treat Equation (6.1) as a fixed effects model, i.e., α_i are treated as fixed for each country as against in a random effects model where α_i are treated as random. Furthermore the fixed effects model imposes restrictions on the parameter β*,* i.e., $\beta_i = \beta$ for all i. This implies that the information on β is pooled. When is pooling appropriate? Only when the homogeneity

[23] Other Pedroni panel statistics were unable to reject the null hypothesis of no co-integration.
[24] Ho,T (2002); 'A panel cointegration approach to saving-investment correlation'; *Empirical Economics*, Vol 27, pp 91-100.
[25] Consider the Equation (5.6) in last chapter. Subtract $\beta_i\Delta x_{it}$ from both sides. Let $u_{it} = (\gamma_{io} - \beta_i)\Delta x_{it} + e_{it}$ then Equation (5.6) reduces to $\Delta(y_{it} - \beta_i x_{it}) = \alpha_i(y_{it-1} - \beta_i x_{it-1}) + u_{it}$. The test of Pedroni tests for unit root in $y_{it} - \beta_i x_{it}$ or equivalently $\alpha_i = 0$ in the reduced equation above. The problem with the residual based test is that it assumes that $\beta_i = \gamma_{i0}$ so that the two errors u_{it} and e_{it} are equal. The violation of this assumption can reduce the power of Pedroni test. To run a checked for the validity of common factor restriction we estimated the individual cross section regression using OLS and compared the estimates of $\beta_i and\ \gamma_{i0}$. We found that except for one country the estimates for other countries varied substantially. Hence this could explain the non-rejection null hypothesis for other Pedroni panel statistics as mentioned in Footnote 23.

assumption, $\beta_i = \beta$ for all i holds. Pesaran and Smith (1995) showed that if homogeneous coefficient is falsely imposed the pooled estimators are inconsistent even if T approaches infinity. However, as pointed by Baltagi (1995, Chapter 4) the pooled model can yield efficient estimates at the expense of a bias, and the researcher must therefore balance the two concerns. Since our interest is to estimate a common relationship across countries we impose the common slope coefficient. Equation (6.1) can now be rewritten as a fixed effects model:

$$w_{it} = \alpha_i + \beta p_{it} + \gamma t + \varepsilon_{it} \tag{6.2}$$

Another reason to proceed with a fixed effects model is that we are interested in a common wage-price relationship for the twelve countries under study. The slope co-efficient captures this as it is the behavioural parameter. We add a trend to maintain consistency with the unit root and the co-integration tests. Equation (6.2) resembles the functional form of bargaining augmented Philips curve in Chapter 4. The long-run relationship can also be specified with price as dependent variable in which case the relationship resembles the wage mark-up relationship. The β is expected close to one and γ should be negative. γ captures the path of real wages across time. One will observe that unlike the equations reported in Chapter 4, our equation does not have a productivity adjustment term. This is because ULC by definition incorporates changes in productivity. Hence the shock to the labour productivity is captured in ULC.

Estimating Equation (6.2) is not straightforward. Kao and Chiang show that OLS estimation may be inconsistent. Following Kao and Chiang's recommendation presented in the previous chapter, we estimate Equation (6.2) using DOLS with one lag and one lead. We tried many combinations of lags and the leads and choose the one above because it minimises the BIC. The following table gives the estimates of β using both OLS and DOLS:

Table 5: Long-run equation: OLS & DOLS

$\mathbf{w}_{it} = \boldsymbol{\alpha}_i + \boldsymbol{\beta}\mathbf{p}_{it} + \boldsymbol{\gamma}\mathbf{t} + \boldsymbol{\varepsilon}_{it}$

	OLS	DOLS
β	1.040 (0.00)	1.057 (0.00)
γ	- 0.007 (0.00)	- 0.007 (0.00)

Note: *Figures in brackets are p-value. F test for H_0: $\gamma = 0$, F(1, 417)= 303.01, p-value:0.00. F test for H_0: β=1, F(1, 417) = 69.56, p-value: 0.00*

Both β and γ are significantly different from zero. The OLS and DOLS estimates do not differ much. We tested the hypothesis H_0: β=1. The F-test rejected the hypothesis at 1%. This implies that on **an average the nominal wage in each country adjust more than fully to price shocks in the long-run**. Since most of the economies included in the panel have active involvement of labour unions in wage determination, the nominal wages adjust more than the price shock because unions' demand nominal wages to rise at least to full extent of the prices and then negotiates on real wage.

The fixed effects from dynamic OLS are presented in Graph 6. Fixed effects capture the unknown country-specific factors. These include factors that may affect wages such as the unemployment rate, union coverage, taxes on wages, producer prices, etc.

Short-run model:

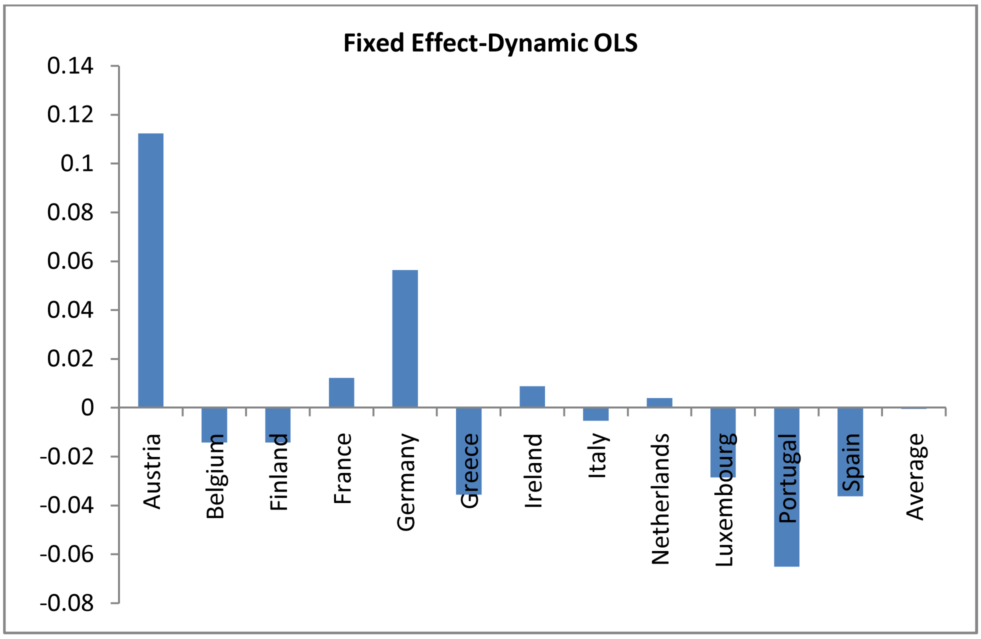
An implication of co-integration is that there must be causality in at least one direction. For this we estimate the following VECM. The VECM is the short-run model and it tells the adjustment mechanism when both wages and prices deviate from the long-run equilibrium in the short-run. We estimate a very simple VECM for the long-run relationship explained in Equation (6.2). The short run equations estimated are as follows:

$$\Delta p_{it} = \alpha_{1i} + \omega_1 \varepsilon_{it\text{-}1} + \sum_{i=1}^{k} \theta_{11i} \Delta p_{it\text{-}i} + \sum_{i=1}^{k} \theta_{12i} \Delta w_{it\text{-}i} + u_{1it} \quad \textbf{(6.3)}$$

$$\Delta w_{it} = \alpha_{2i} + \omega_2 \varepsilon_{it\text{-}1} + \sum_{i=1}^{k} \theta_{21i} \Delta p_{it\text{-}i} + \sum_{i=1}^{k} \theta_{22i} \Delta w_{it\text{-}i} + u_{2it} \quad \textbf{(6.4)}$$

The errors for period t-1 are estimated from the long-run Equation (6.2) after imposing the DOLS estimates for fixed effects, β and γ on Equation (6.2). The coefficients ω_1 and ω_2 are the adjustment parameters and they tell the degree to which of the respective left side variables adjust in period t to disequilibrium shocks in period t-1.[26] The sign for ω_1 should be positive and that of ω_2 negative to ensure that equilibrium is restored in the long-run.

Graph 6



Equations (6.3) and (6.4) are dynamic as they involve lags of the dependent variables. Furthermore, both the equations have fixed effects hence the OLS estimate will be biased and inconsistent. As explained in Chapter 5 estimating Equations (6.3) and (6.4) in first

[26] Readers may find it curious that Westerlund's test only uses one error correction equation to test for the co-integration while we have two error correction equations in our short-run model. The intuition for this is that the test assumes x_{it} is weakly exogeneous with respect to α_i and β_i and under this assumption, α_i and β_i can be efficiently estimated using Equation (5.6). Thus weak exogenity implies that x_{it} is not error correcting. The assumption enables us to implement the test using Equation (5.6) but does not preclude the possibility that x_{it} can also be error correcting. Furthermore, α_i in Equation (5.6) is shown to be related to ω_1 and ω_2 in Equations (6.3) and (6.4) by this relation:
$\alpha_i = \omega_{iy} - \gamma_{i0}\omega_{ix}$. However if weak exogenity fails then using a test based on Equation (5.6) will be erronous. [See Section 2.2 of Westurlund (2007) for further details.] In Westerlund's test weak exogeneity is a maintained assumption.

differences has endogeneity because the right hand side term is correlated with the differenced error term. In such cases the standard procedure involves using instrumental variable technique or the generalised method of moments.[27] Following our discussion in Chapter 5 for dynamic panel models, we see that we have two options to estimate the Equations (6.3) and (6.4). First, the LSDV estimate by following the recommendation in Table 2. LSDV is also appropriate because the average number of time observation in the panel is 39 (>30). Hence the Nickel bias is expected to be small. Second option is Bruno (2005) proposed estimator by correcting the bias in LSDV estimator for unbalanced panel data with N small relative to T. Bruno (2005) requires specifying an initial estimator to approximate the bias. We use GMM1 and AH for our case by using recommendation from Table 2. Since Bruno (2005) used a stylised panel of (N=10, T =40) for their Monte Carlo study, which is very close to our panel dimensions (N=12, T =39), it will be instructive to see the relative difference from the three estimators. The estimates are presented in Tables 6 to 8.

We summarise the findings from our short-run estimations as under using table 8:

1. The nominal wages adjust faster than price to a disequilibrium shock from the previous period. A similar conclusion for the Dutch economy was established by Broersma and Butter (1999) where they conclude that 'the Dutch labour market is known to adjust relatively quickly to adverse shocks'. Fountas, Lally and Wu (1999) also report in their study for Ireland that nominal wages adjust faster than price. The adjustment coefficients in their study for Ireland were 0.123 for wages and -0.018 for prices (signs are reversed as their long-run equation was specified in terms of prices). We can conclude that labour markets adjust faster on an average than goods market in all twelve countries in our panel.

Table 6: Short-run equations (pooled) using LSDV

	ε_{it-1}	Δp_{it-1}	Δp_{it-2}	Δw_{it-1}	Δw_{it-2}
Δp_{it}	0.063 (0.00)	0.911 (0.00)	-0.279 (0.00)	0.089 (0.02)	0.127 (0.00)
Δw_{it}	-0.111 (0.00)	0.728 (0.00)	-0.341 (0.00)	0.519 (0.00)	-

Note: *Figures in brackets are p-values. The appropriate lag lengths selected using BIC criterion.*

[27] Costantini and Martini (2009) in estimating the VECM for the long-run relation between energy consumption and economic growth use the Arellano and Bond (1991) GMM1 estimator. Damette and Froute (2010) in their panel study for the long-run relation between risk aversion and CDS spreads also use the restricted GMM1 with two lags of dependent variable as instruments.

Table 7: Short-run equations (pooled) using LSDV Bias-corrected

(Bias-correction using Anderson-Hsiao Instrumental variable estimator)

	ε_{it-1}	Δp_{it-1}	Δp_{it-2}	Δw_{it-1}	Δw_{it-2}
Δp_{it}	0.063 (0.28)	0.911 (0.00)	-0.279 (0.02)	0.089 (0.32)	0.127 (0.20)
Δw_{it}	-0.091 (0.16)	0.734 (0.00)	-0.293 (0.09)	0.537 (0.00)	-

Note: *Figures in brackets are p-values. Bias correction initialised by Anderson and Hsiao estimator using 50 bootstrap replications to obtain the robust standard errors and the variance-covariance matrix. Bias corrected up to O(1/NT. The appropriate lag lengths are selected using BIC criterion. For the price equation: AH instrumented* Δp_{it-1} *using the second lag of* Δp_{it} *and first differences of* ε_{it-1}, Δp_{it-2}, Δw_{it-1}, Δw_{it-2}. *For the wage equation: AH instrumented* Δw_{it-1} *using the second lag of* Δw_{it} *and the first differences of* ε_{it-1}, Δp_{it-2}, Δw_{it-1}, Δw_{it-2}.

Table 8: Short-run equations (pooled) using LSDV Bias-corrected

(Bias-correction using Arellano-Bond GMM estimator)

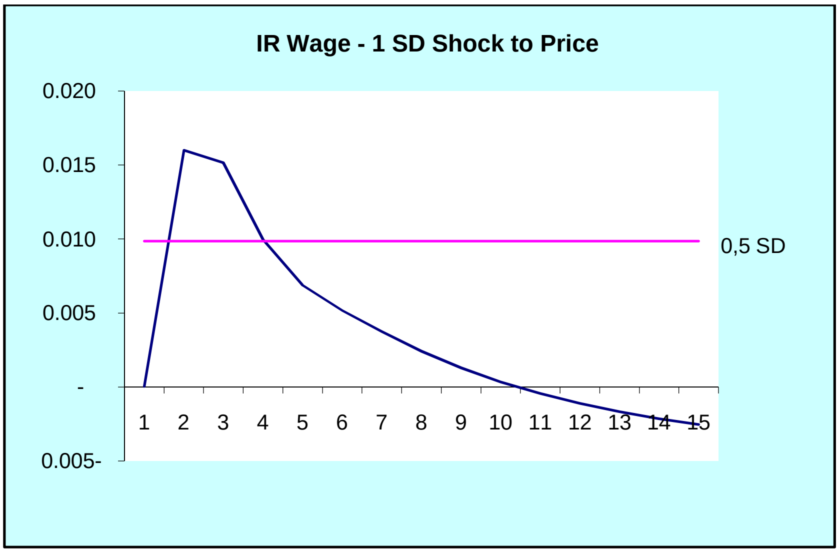
	ε_{it-1}	Δp_{it-1}	Δp_{it-2}	Δw_{it-1}	Δw_{it-2}
Δp_{it}	0.044 (0.04)	0.967 (0.00)	-0.299 (0.00)	0.096 (0.04)	0.137 (0.00)
Δw_{it}	-0.103 (0.00)	0.729 (0.00)	-0.371 (0.00)	0.559 (0.00)	-
Wald's Test		Price equation H_0: $\Delta w_{it-1} = \Delta w_{it-2} = 0$ $\chi 2_{(2)} = 28.68$ (0.00)		Wage equation H_0: $\Delta p_{it-1} = \Delta p_{it-2} = 0$ $\chi 2_{(2)} = 151.80$ (0.00)	

Note: *Figures in brackets are p-values. Bias correction initialised by Arellano and Bond estimator using 50 bootstrap replications to obtain robust standard errors and the variance-covariance matrix. Bias corrected up to O(1/NT).The appropriate lag lengths are selected using BIC criterion. For the price equation: AB instrumented* Δp_{it-1} *using the lags of* Δp_{it}, *and first difference of* ε_{it-1}, Δp_{it-2}, Δw_{it-1}, Δw_{it-2}. *For the wage equation: AB instrumented* Δw_{it-1} *using the lags of* Δw_{it}, *and the first differences of* ε_{it-1}, Δp_{it-2}, Δw_{it-1}, Δw_{it-2} *[See Footnote 19, page 25 for an explanation.]*

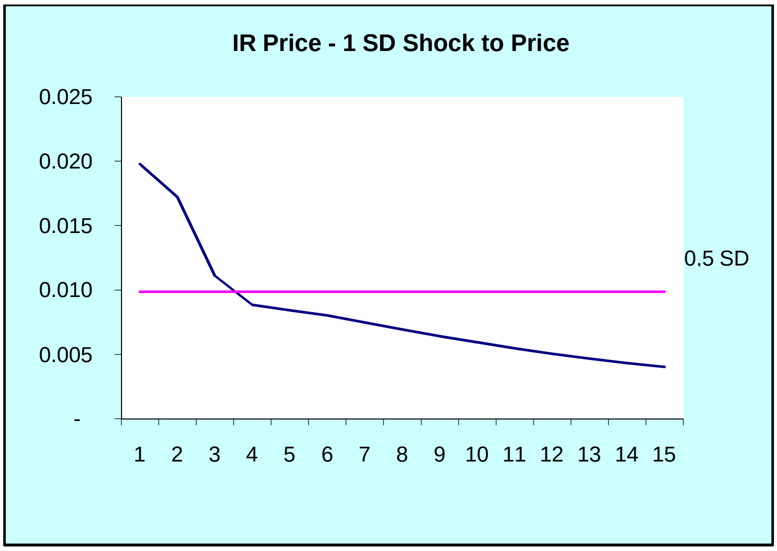
2. The impulse response functions for one-standard deviation shock to price (Graphs 7 and 8) and one-standard deviation shock to wage (Graphs 9 and 10) are shown below. A shock to price at time zero immediately translates to wages. Then over time both variables revert to equilibrium. Similarly for a 1SD shock to wage translates into less than 0.5 SD reactions to price.

3. The change in nominal wages in current period depends on last period's price inflation (coefficient 0.729) more than changes in nominal wages in the last period (coefficient 0.559). Similarly the change in price inflation in the current period depends on changes inflation in the last period (coefficient 0.967). Furthermore, a rise in the nominal wage in the last period only effects inflation in current period marginally (coefficient 0.096). Hence the causality that price inflation causes wage inflation is stronger that wages inflation causing price inflation. Hence, the bargaining augmented wage equation better explains the wage-price relationship in the twelve countries than the mark-up theory. The evidence for the mark-up relationship is weak in general.

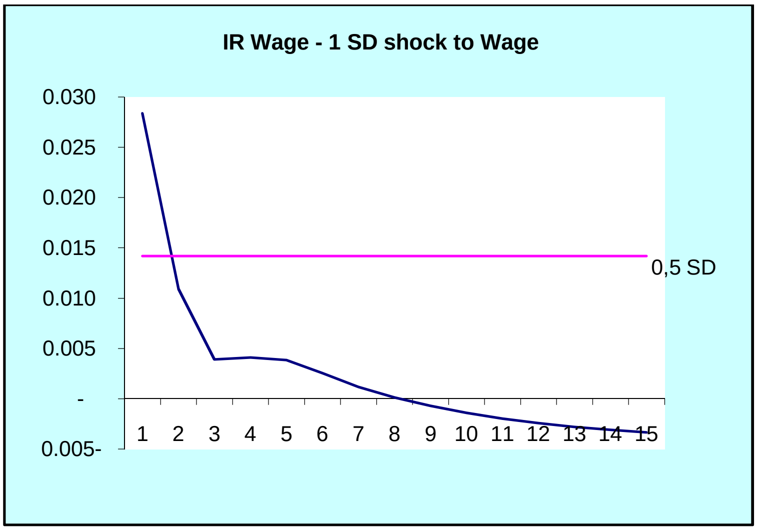
Graph 7



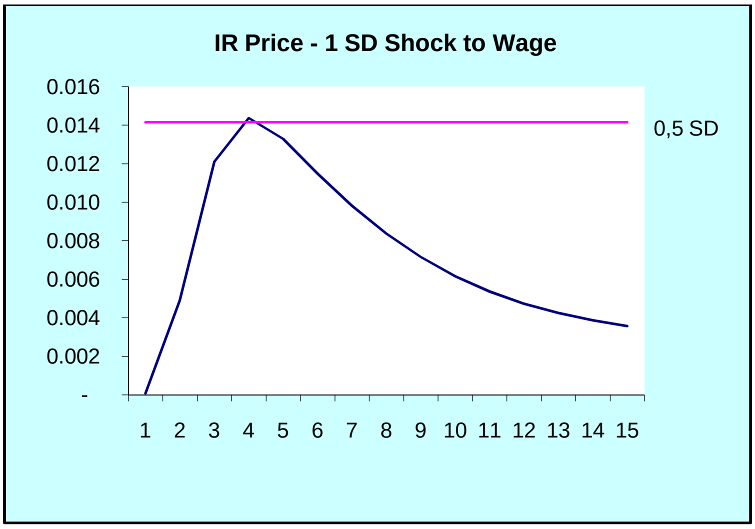
Graph 8



Graph 9



Graph 10



VII

Conclusions and way forward

1. The empirical exercise suggests that wages and prices are co-integrated and that wages adjust to prices by a factor slightly more than one in the long-run. That is, the estimated hedge ratio is greater than one. Given the assumptions- (i) that the emergence of HICP does not alter the behaviour of economic agents, and (ii) the long-run co-integration equation assumes a trend, this implies that for pension funds that index their pension liabilities to wage inflation, derivatives based on HICP are effective hedges in the long-run. Hedging is effective because co-integration between wages and prices ensure that the basis risk is minimised in the long-run even though wage inflation diverges from price inflation in the short-run. Alternatively, we can also say that, although the risk of wage inflation and the risk of price inflation appear as two separate risks in the short-run, in the long-run they are related, and therefore one risk is fully expressed in terms of the other. Hence, the use of HICP derivatives to hedge the risk of indexation will have minimal basis risk in the long-run given the convergence of national CPIs and national HICPs.

2. As a corollary to the previous argument, it is unadvisable to hedge the risk of indexation on a short-run basis using inflation linked instruments because the hedge quality is poor due large deviations between wage inflation and price inflation in the short-term. Hence it is recommended that pension funds hedge the risk of indexation using HICP only on a long run basis.

3. Co-integration analysis of pension fund risks like the one presented in this thesis has profound impact on asset-liability management (ALM). The current practice in ALM is to view the long-run as a series of short periods. This is achieved by generating economic scenarios (typically using VAR models) for one time step in the future till the desired time horizon. This approach therefore assumes that a single long-run decision is equivalent to a series of short-run decisions. However co-integration views the long-run as one single time block. This difference can lead to very different outcomes. This is because a single long decision is not the same as series of short decisions. Secondly, the long-run dynamics explained by VAR will represent the true long-run dynamics only when the long-run error is explicitly accounted for in the VAR model. In ALM models co-integration is generally ignored, and therefore the second-order feedback from long-run errors is missing. Hence there will be a departure from the true long-run outcome. Furthermore, pension funds are perceived as long-run investors. Hence in light of the above arguments, current ALM practices should be modified so that they address the long-run decision making characteristics of pension funds.

4. If a pension fund wants to buy a cost efficient hedge for wage inflation, lack of markets for such a risk implies paying high premium for the insurance. The co-integration framework as suggested here helps in finding cost effective hedging instruments like inflation linked derivative for the risk of wage inflation. This is because when wage and prices are co-integrated correlation between wage and prices will increase over time and tend to one. Thus the hedge effectiveness of inflation linked derivative improves over time. The liquid markets for HICP derivatives make hedging cost efficient.

Criticisms and improvements: The current study may be criticised on following points:

1. First criticism could be that this study models the wage price mechanism for total economy, but for the Netherlands where pension funds are organised on sectors, sector specific adjustments can be different. Yes, this is true and we can extend this by including sector specific variables for wages. But availability of data can be an issue here.

2. It can be argued that this model is based in closed economy framework while theory suggests that domestic inflation is sensitive to world prices for small open economy. Yes, this point is valid but it only changes how we define the measure of prices. One way to include imports is to measure prices as a weighted average of domestic prices and world prices.

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Appendix – A

A Short note on definition of co-integration

Nelson and Plosser (1982) introduced the idea that economic series have stochastic trend. In a univariate time series case stochastic trend can be removed by taking the first difference. It is tempting to generalise this idea of differencing to a multivariate case. However it is not appropriate to treat non-stationary variable by simply taking first difference in multivariate case because it is quite possible that there will be a linear combinations of non-stationary variables (that is series with unit roots) that may be stationary. This idea is referred to as co-integration and such variables are referred to as the co-integrated variables. The presence of co-integration and the resultant error correction mechanism is central to understanding the short-long dynamics of variables in question. Formally, co-integration is defined as follows:

[Engle and Granger, (1987] A set of economic variable are in long run equilibrium when:
$\beta_1 x_{1t} + \beta_2 x_{2t} + \ldots + \beta_n x_{nt} = \mathbf{0}$
Let $\boldsymbol{\beta} = (\beta_1, \beta_2, \beta_3, \ldots \beta_n)$ *and* $\mathbf{X_t} = (x_{1t}, x_{2t}, x_{3t}, \ldots x_{nt})^T$ *then deviation from long run equilibrium – called the equilibrium error – denoted by* e_t *so that* $e_t = \boldsymbol{\beta}\mathbf{X_t}$
If the equilibrium is meaningful, then the equilibrium error must be a stationary process. Now the components of the vector $\mathbf{X_t}$ *are co-integrated of order (d, b) if –*

1) *All components of* $\mathbf{X_t}$ *are integrated of order d*
2) *there exists a vector* $\boldsymbol{\beta} = (\beta_1, \beta_2, \beta_3, \ldots \beta_n)$ *such that the linear combination* $\boldsymbol{\beta}\mathbf{X_t} = \beta_1 x_{1t} + \beta_2 x_{2t} + \ldots + \beta_n x_{nt}$ *is integrated of order (d-b), b>0. The vector* $\boldsymbol{\beta}$ *is called the co-integrating vector.*

The term equilibrium refers to statistical notion of the long-run relationship between the non-stationary variables as against the economic theory idea of the desired and the actual outcome. Furthermore, co-integration does not require that the long-run relationship be generated by market forces or by behavioural rules of the agents. As per Engle and Granger (1987) the equilibrium relationship may be a causal behaviour, or simply a reduced-form relationship among similar trending variables.

A key feature of the co-integrating variables is that their evolution in time is determined by the deviation from the long-run relationship. If a system of variables were to return to the long-run equilibrium then at least some variables must respond to the disequilibrium. Further there are empty numbers of ways in which the variables may adjust to regain the long-run relationship. To identify which of the adjustment mechanism is relevant among the many possible choices is described by an error correction mechanism. Formally, the error correction model is defined as:

The vector $\boldsymbol{X_t} = (x_{1t}, x_{2t}, x_{3t} \ldots x_{nt})^T$ *has an error correction representation if it can be expressed in the form:*

$$\Delta X_t = \Pi_0 + \Pi X_{t-1} + \Pi_1 \Delta X_{t-1} + \Pi_2 \Delta X_{t-2} + \ldots + \Pi_p \Delta X_{t-p} + \varepsilon_t$$

where,

$\boldsymbol{\Pi_0}$ *= (n x 1) vector of intercept terms,*

$\boldsymbol{\Pi_i}$ *= (n x n) coefficient matrix*

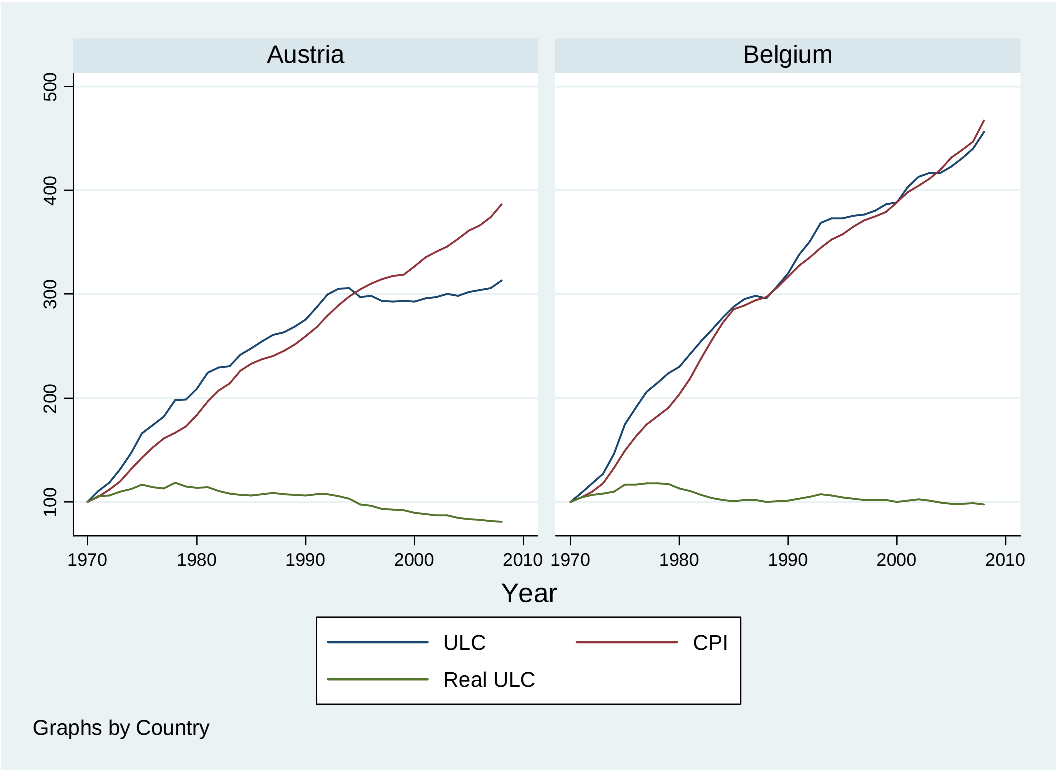
$\boldsymbol{\Pi}$ *= matrix with elements such that one or more* $\Pi_{jk} \neq 0$

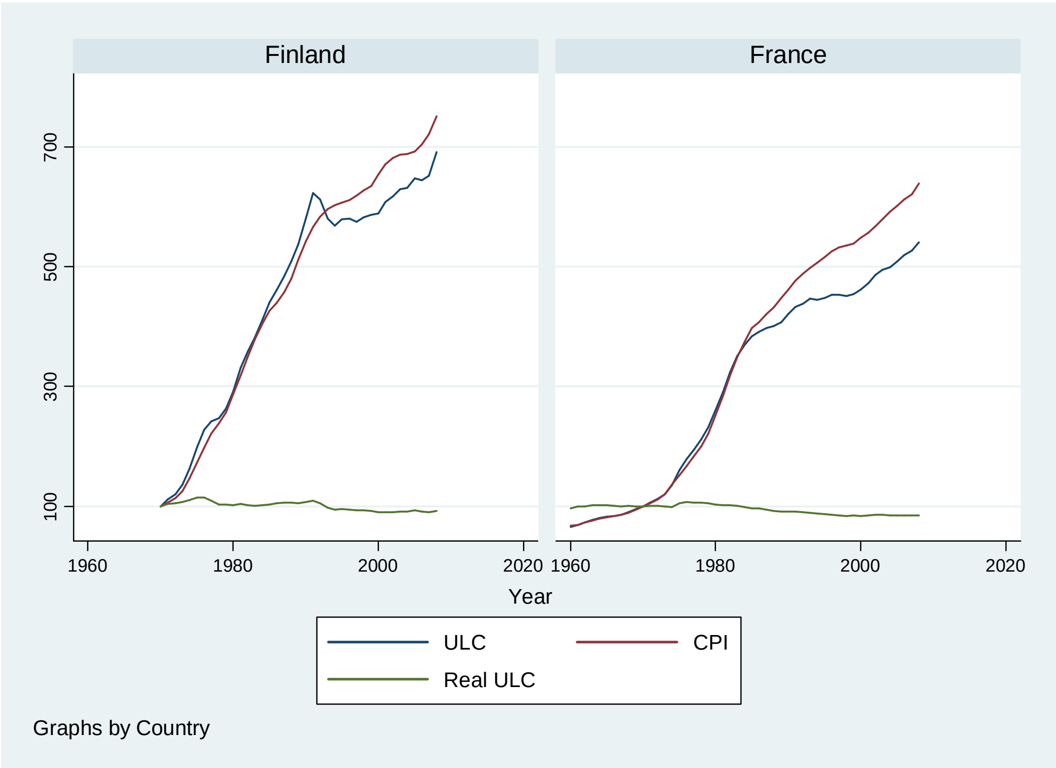
$\boldsymbol{\varepsilon_t}$ *= vector of error terms (such that* $\boldsymbol{\varepsilon_{it}}$ *may be correlated to* $\boldsymbol{\varepsilon_{jt}}$*)*

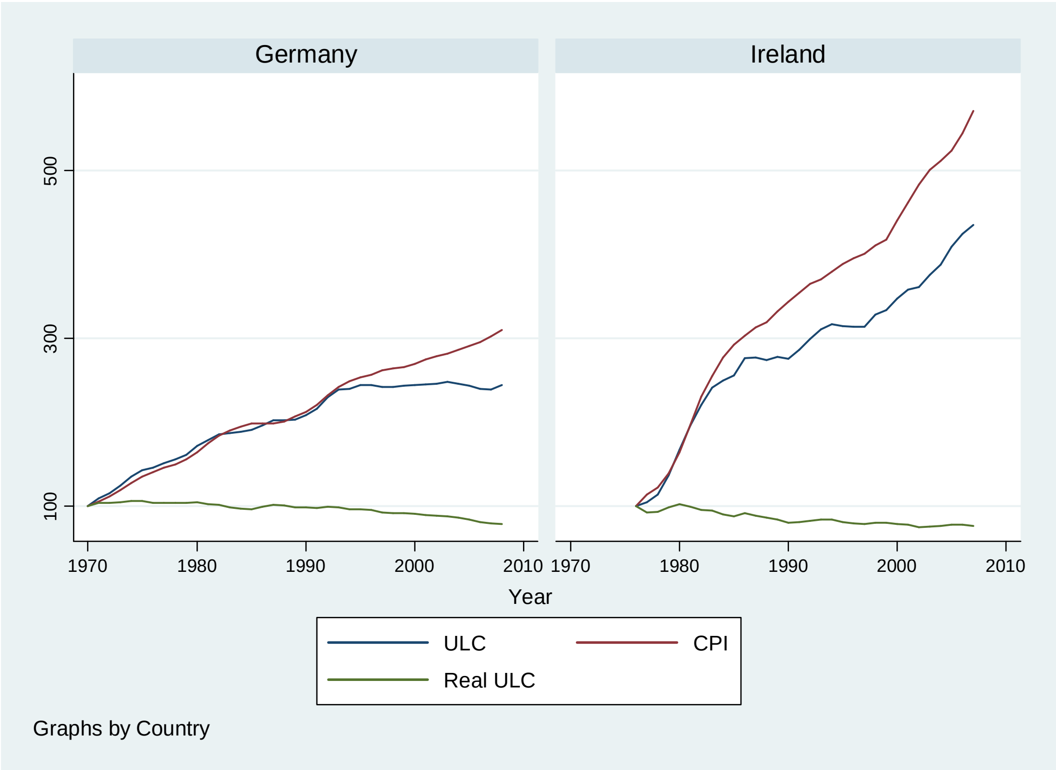
The starting point for a co-integration analysis is therefore to ascertain whether the underlying series are non-stationary that is there is a unit root. We then test for the presence of co-integration among non-stationary variables. Monte Carlo simulations on unit root and co-integration tests show that time series tests have limited power. Further in empirical research availability of long time series constraints the use of co-integration analysis in a time series framework. One way to augment the power of unit roots test is to increase the sample size. This can be achieved by using a pooled (panel) data. The underling theory for such an approach is the application of the following principle: ‘if we have n independent and unbiased estimates of a parameter then mean of the estimates is also unbiased. Further so long as the estimates are independent, by central limit theorem; the mean will be normally distributed’.

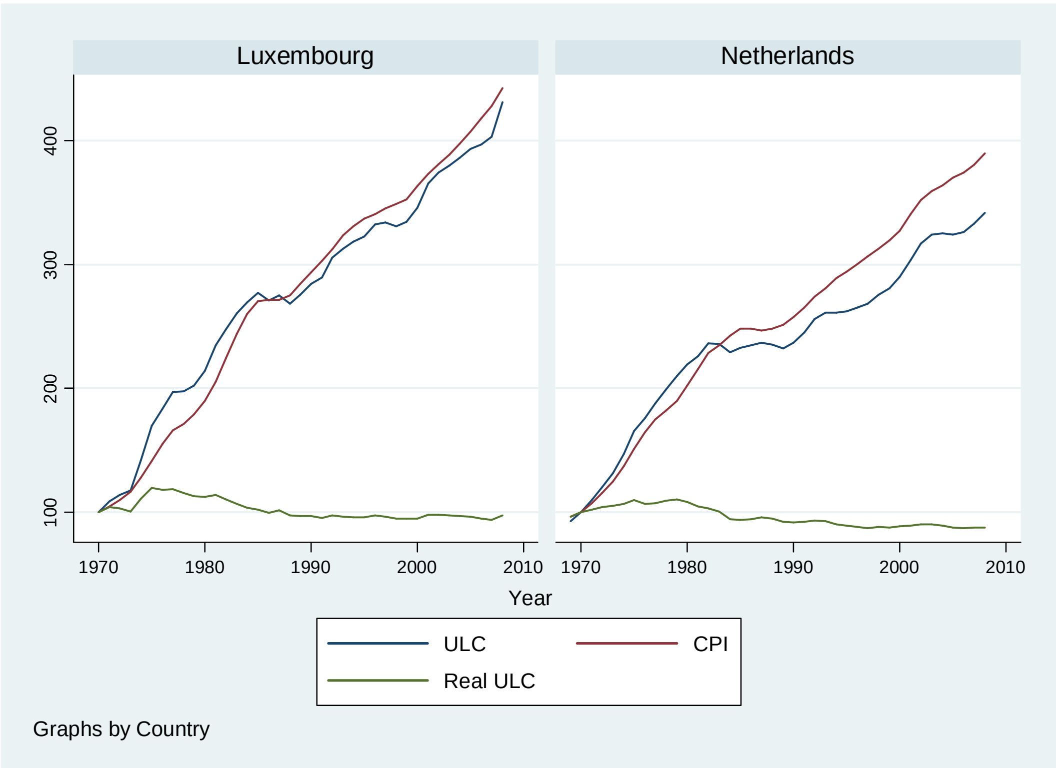
Appendix B

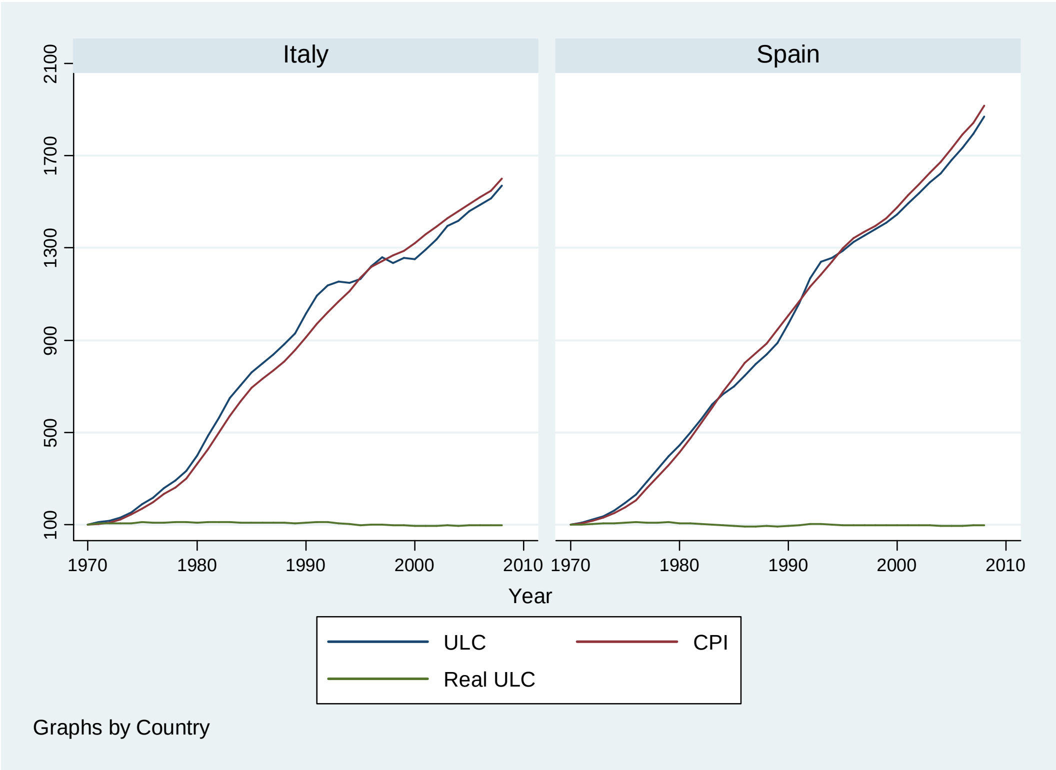
Country-by-country plot of the raw-data

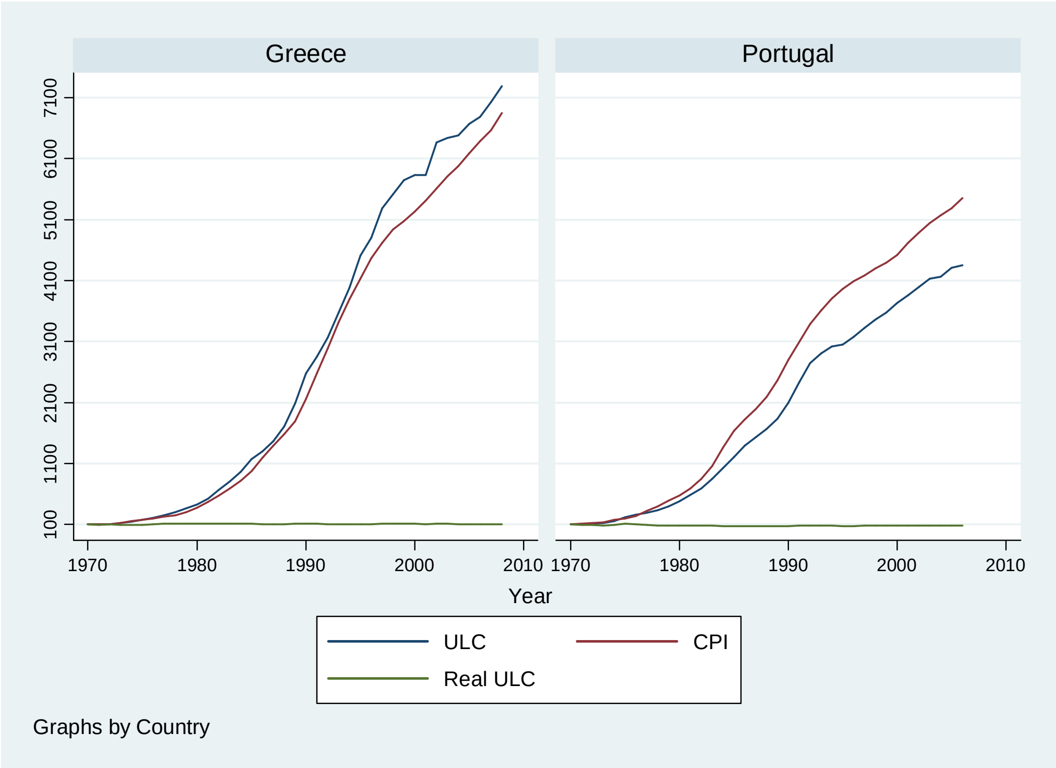





Germany
Ireland
100
300
500
1970
1980
1990
2000
2010
Year
ULC
CPI
Real ULC
Graphs by Country


Luxembourg
Netherlands
100
200
300
400
1970
1980
1990
2000
2010
Year
ULC
CPI
Real ULC
Graphs by Country


Italy
Spain
2100
1700
1300
900
500
100
1970
1980
1990
2000
2010
Year
ULC
CPI
Real ULC
Graphs by Country


Greece
Portugal
7100
6100
5100
4100
3100
2100
1100
100
1970
1980
1990
2000
2010
Year
ULC
CPI
Real ULC
Graphs by Country

Appendix C

Country-by-country plot of log-first differences