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Abstract

Cost-of-Living-Adjustment (COLA) coverage figures suggest a time variation in the degree of wage indexation. In spite of this observation, most current literature conveniently assume a constant degree of indexation as this variable is not directly observable. This study intends to empirically measure the time variation in the degree of wage indexation. To this end, we derive a reduced form version of the New Keynesian Wage Phillips Curve under the assumption of a time varying degree of wage indexation. A state-space methodology is then employed in estimating this model using data of selected OECD countries. The study subsequently investigates variables influencing the time variation in the degree of wage indexation. Our results consistently suggest a substantial time variation in the degree of wage indexation in all countries considered. The wage indexation estimates obtained for the US bear remarkable similarities with the figures suggested by COLA coverage. It is subsequently shown that variations in trend inflation significantly explain variations in the degree of wage indexation. Finally, there is weak evidence in support of the Gray hypothesis that wage indexation is negatively correlated with the variance of productivity shocks.

JEL Classification Codes: E12, E24, E31, E40

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1 Introduction

The proportion of wage contracts with cost-of-living-adjustment (COLA) clauses in the US has been observed to vary between 20% and 62% since the 1950s. This can be construed as evidence for time variation in the degree of wage indexation since this percentage of COLA coverage is a widely accepted proxy for the degree of wage indexation in the US.[1] In spite of this evidence, a substantial proportion of theoretical research on the topic typically assumes a constant degree of wage indexation. Furthermore, some recent empirical studies devoted to estimating the degree of wage indexation give a time invariant estimate to this parameter.[2]

The COLA coverage figures suggest that models regarding wage indexation should incorporate the time-varying nature of wage indexation. In addition to this, there are other motivations as to why wage indexation models should consider incorporating the time-varying nature of wage indexation. We outline three of such motivations in the subsequent paragraphs.

First, the time-varying nature of wage indexation has implications for the unconditional distributions of macroeconomic variables. Current models work under the assumption of a constant degree of wage indexation. A consequence of this assumption is that macroeconomic variables are normally distributed. However, empirical evidence as documented in Chang (2012), for instance, supports the existence of fat tails in the distribution of inflation. Attey and de Vries (2011) provides a possible theoretical explanation for this empirical observation. This explanation is linked to the time-varying nature of wage indexation. The aforementioned study derives a new Classical Phillips curve under the assumption of random wage indexation and solves for equilibrium inflation in a version of the Barro-Gordon model. It is subsequently shown that under this model, an unconditional distribution of inflation exists and is fat tailed. The intuition behind this result is that shocks to the degree of wage indexation may act as multiplicative shocks rather than additive shocks. Therefore, these multiplicative shocks exacerbate the effects of any extreme realizations of other (additive) shocks to inflation, thus producing the fat tail.

Second, incorporating time variation in the degree of wage indexation into models enables one to gain additional insights into factors explaining the volatility of macroeconomic variables. For instance, it is conceivable that the distribution of the degree wage of indexation is determined ,at least in part, by labour market institutional variables such as bargaining power of unions. One can exploit this link to investigate the relationship between labour market institutional factors and the volatility of inflation since the latter variable depends on the distribution of the time-varying degree of wage indexation. Attey and Kouame (2015) confirm that the correlations between inflation volatility and labour market institutional variables are often significant. This correlation would not have been obvious if wage indexation models abstract from the time-varying nature of wage indexation.

Finally, the preceding two motivations imply that the presence of a time variation in the degree of wage indexation does have some implications for the conduct of monetary policy. A monetary policy conducted by a Taylor rule, for instance, necessitates the response of the interest rate to shocks stemming from the

[1] Figure 5a contains a time plot of COLA coverage in the US for the period spanning 1955 to 1995 after which it was discontinued

[2] Smets and Wouters (2003) estimate a dynamic stochastic general equilibrium (DSGE) model for the Euro area. The estimate for the degree of wage indexation has a posterior mean of 0.728. More recently, Gali (2011) and Muto and Shintani (2014) estimate a New Keynesian Wage Phillips Curve for the US and Japan respectively. The estimates for the degree of wage indexation in both are statistically significant estimates.

degree of wage indexation. Also, the determinacy of a system associated with a policy rule depends on the degree of wage indexation. Ascari et al. (2011) show how the probability of determinacy of a system characterized by a Taylor rule depends on the level of wage indexation. In particular, the study finds that a higher degree of wage indexation increases the probability of a system being determinate. Thus, the response of macroeconomic indicators to monetary policy in the US during the period spanning the mid-1970s to early 1980s when wage indexation was relatively higher might differ from the response in other periods when wage indexation was lower, given the same Taylor rule parameters.

It stands to reason that the descriptive and prescriptive performances of wage indexation models would be greatly improved by incorporating a time-varying degree of wage indexation. However the unobservable nature of this variable limits the accuracy of this class of models. The main purpose of this paper is to develop an estimation methodology for the time-varying degree of indexation. The model employed in this study augments that of Gali (2011) with a time-varying degree of indexation and productivity growth. The resulting reduced form expression which is labeled the time-varying New Keynesian Wage Phillips Curve (TV-NKWPC) is estimated using a state-space methodology. This methodology permits one to capture the time variations in the degree of wage indexation. The estimation is done using data of US and 10 other OECD countries, namely: Austria, Belgium, Canada, Finland, Germany, Japan, Netherlands, Norway, Sweden and the UK.

This study further investigates the factors that explain the time variation in wage indexation. In order to do this, the study performs country specific OLS estimations of wage indexation equations with trend inflation, (time-varying) variance of productivity shocks and other labour market institutional variables as explanatory variables. Given the findings in numerous studies indicating a positive relationship between inflation uncertainty (which is positively correlated with trend inflation) and wage indexation, it is expected that trend inflation will have a significant effect.

The variance of productivity shocks is included in the list of regressors in order to have an ad hoc test of the empirical validity of a hypothesis derived from Gray (1976), which predicts a negative relationship between the degree of wage indexation and productivity shocks. The test might best be described as ad hoc since the original hypothesis relies on the assumption that wage indexation is a policy instrument used by a policy maker. This study makes no such assumption. The labour market institutional variables are included to control for the bargaining power of unions and other variables that might explain the time variation in wage indexation.

Our study is not the only attempt at estimating time variations in the degree of wage indexation. Ascari et al. (2011) and Hofmann et al. (2010) have also attempted to estimate the time variations in the degree of wage indexation. The former study employs a methodology based on a rolling-window OLS regression of wage inflation on its lags and lags of price inflation. The latter study adopts a methodology based on a Bayesian VAR approach with time-varying coefficients. The estimates of the time-varying degree of wage indexation in the two studies are consistent with the general belief that wage indexation continuously fell during and after the great moderation. However, the specific values of the estimates do sometimes deviate from wage indexation figures suggested by the proportion of COLA covered contracts. For instance, estimates provided by Ascari et al. (2011) peaked around 0.9 during the 'Great Inflation' period while COLA figures suggest 0.61 as the highest value for wage indexation during this period. Also, the estimates by Hofmann et al. (2010) are less than 0.5 throughout the whole sample period. Furthermore, the methodology adopted in this paper is simpler than those of Ascari et al. (2011) and Hofmann et al. (2010).

The remainder of this paper is organized as follows. The time-varying New Keynesian Phillips Curve is derived in Section 2, where the NKWPC is shown to be a special case of the more general TV-NKWPC. The first part of Section 3 includes some diagnostic tests on an estimated NKWPC using US data in order to provide evidence for the presence of time variation in the degree of wage indexation. The second part develops and estimates the state-space regression model of the TV-NKWPC. Section 4 provides country-specific estimates of the TV-NKWPC for 11 OECD countries and also estimates for the OLS regression of the wage indexation equation. Finally, Section 5 concludes.

2 The New Keynesian Wage Phillips Curve

Gali (2011) derives a New Keynesian Wage Phillips Curve (NKWPC) based on the assumption of staggered wage setting by the representative household. We extend this model by incorporating a time-varying degree of wage indexation. The resulting expression is designated as the time-varying New Keynesian Wage Phillips Curve (TV-NKWPC). This section briefly explains the theoretical derivation of the TV-NKWPC and shows how the NKWPC is a special case of the more general TV-NKWPC.

Consider a representative household with members who can be represented by the unit square and indexed by a pair $(i, j) \in [0, 1] \times [0, 1]$, with the first dimension i representing labour type and the second dimension determining their disutility from work. Let the disutility from supplying labour type j be $\chi_t j^{\varphi}$ where the variable χ_t denotes the exogenous labour supply shock. Assume that consumption (C_t) enters utility function in a loglinear manner. This implies the following expression for the utility function:

$$\begin{aligned} U(C_t, N_t(i), \chi_t) &= \log C_t - \chi_t \int_0^1 \int_0^{N_t(i)} j^{\varphi} dj \ \ di \\ &= \log C_t - \chi_t \int_0^1 \frac{N_t(i)^{1+\varphi}}{1+\varphi} di. \end{aligned}$$

Further assume each household member supplies specialized labour which is an imperfect substitute to other members' labour supply. The aggregate labour index by the household has the following Dixit-Stiglitz form:

$$N_t \equiv \left[\int_0^1 N_t(i)^{1-\frac{1}{\epsilon}} di \right]^{\frac{\epsilon}{\epsilon-1}},$$

whereby ϵ denotes the elasticity of substitution between the different labour types. An intratemporal problem of cost minimization given a wage rate $W_t(i)$ by the members of the household yields the following expression for labour supply of type i:

$$N_t(i) = \left(\frac{W_t(i)}{W_t} \right)^{-\epsilon} N_t.$$

The variable W_t denotes the aggregate wage index with its expression implicitly given as follows:

$$W_t \equiv \left[\int_0^1 W_t(i)^{1-\epsilon} di \right]^{\frac{1}{1-\epsilon}}$$

The representative household seeks to maximize its lifetime utility subject to its budget constraint. The objective function of the household and the budget constraint are respectively given below:

$$E_0 \sum_{t=0}^{\infty} \beta^t U(C_t, N_t(i)) \tag{1}$$

$$P_t C_t + Q_t B_t \leq B_{t-1} + \int_0^1 W_t(i) N_t(i) di + \Upsilon_t. \tag{2}$$

The variable P_t represents the price level while B_t represents one-period riskless bond purchased at price Q_t. The variable Υ_t denotes the lump-sum component of income. The constraint in (2) is supplemented by the usual transversality conditions to prevent bubble solutions.

2.1 Time-varying wage indexation

In each period, a worker resets their nominal wage with probability $1 - \theta$. Workers who do not get the opportunity to reset have their wages automatically indexed according to the following indexation rule:

$$W_{t+k|t} = W_t^* X_{t+k|t}, \tag{3}$$

where W_t^* is the optimal nominal wage level prevailing at time t for a worker who resets their wages in that period. The X_{t+k} is generally a function of inflation and other variables to which wages are indexed. Similar to the indexation rule found in Fischer (1983) and Jadresic (2002), it is assumed that workers index to both productivity and inflation.[3] Let $X_{t+k|t} = \exp(x_{t+k|t})$. The following expression for log indexation $(x_{t+k|t})$ is proposed:

$$x_{t+k|t} = \begin{cases} 0 & k = 0 \\ \sum_{s=0}^{k-1} \left(\gamma_{t+s+1}\bar{\pi}^p_{t+s} + (1-\gamma_{t+s+1})\pi^p + \phi\pi^z_{t+s+1} + (1-\phi)\pi^z\right) & k \geq 1. \end{cases} \tag{4}$$

where $\bar{\pi}^p_t$ and π^p denote the inflation rate (or its moving average) implied by the indexation agreement and the steady-state inflation rate respectively. The variables π^z_t and π^z denote the growth in productivity and its steady-state value respectively.

While the general features of the wage indexation rules found in the literature allow for log wages (w_t) to react in a deterministic manner to an inflation measure ($\bar{\pi}^p_t$) and productivity growth (π^z_t), our indexation rule additionally allows for the possibility of time variation in the degree of wage indexation to inflation, γ_t. Empirical estimates such as those found in Holland (1986), Ascari et al. (2011) can be interpreted as evidence for the time-varying nature of wage indexation to inflation. For this reason, we time index γ while assuming that the influence of productivity growth on the indexed part of wages is time invariant.[4] The variation in wage indexation might reflect, for instance, the varying bargaining power of unions. Also, a time-varying γ_t is more compatible with the observation that wage indexation is higher in the presence of a higher level of (trend) inflation.

[3]While the indexation rules used in the literature cited imply that wages are indexed to output and inflation, it is assumed here that wages are indexed to productivity instead of output.

[4]While this assumption may seem arbitrary, estimations provided in Table 12 do not reject this assumption.

As will be shown later in this section, the New Keynesian Wage Phillips Curve (NKWPC) derived under the assumption of time-varying degree of wage indexation exhibits time-varying parameters. The expression for aggregate wages (W_t) implied by the indexation expression (4) is given as follows:

$$W_t = \left[\theta \left(W_{t-1}X_{t|t-1}\right)^{1-\epsilon} + (1-\theta)(W_t^*)^{1-\epsilon}\right]^{\frac{1}{1-\epsilon}}. \tag{5}$$

2.2 Staggered wage setting and the NKWPC

Similar to the wage setting mechanism in Erceg et al. (2000), it is assumed that a worker resets nominal wages with probability $1-\theta$. A worker that resets their wages in period t chooses nominal wages to maximize their lifetime utility given by the equation (1) subject to the constraint implied by the demand for their labour. The first order condition for the household is given as follows:

$$\sum_{k=0}^{\infty}(\beta\theta)^k E_t \left[N_{t+k|t}U_c(\cdot)\left(\frac{W_t^* X_{t+k|t}}{P_{t+k}} - \mathcal{M}MRS_{t+k|t}\right)\right] = 0, \tag{6}$$

where MRS_t denotes the marginal rate of substitution between consumption and labour, $\mathcal{M} \equiv \epsilon/(\epsilon-1)$ denotes the wage mark-up under flexible prices and β denotes the discount factor. The specification of the utility function implies that the expression for marginal rate of substitution can be derived as follows: $MRS_{t+k|t} = C_{t+k}N^{\varphi}_{t+k|t}$. We loglinearize the expression (5) around a deterministic steady state. Substituting the resulting expression as well as (4) into a loglinearized version of (6) results in the following expression:

$$\pi_t^w - \nu_t = \beta E_t(\pi_{t+1}^w - \nu_{t+1}) - \lambda(\mu_t - \mu), \tag{7}$$

where $\nu_t = x_{t|t-1}$ and $\lambda = [(1-\theta)(1-\beta\theta)]/[(1+\epsilon\varphi)\theta]$. The variable π_t^w indicates the growth rate (defined as log-difference) of wages. The variable μ_t denotes the average markup defined as the difference between the log of real wages and the marginal rate of substitution. The expression of μ_t is given below:

$$\mu_t = w_t - p_t - [c_t + \varphi n_t + \log(\chi_t)]. \tag{8}$$

In the flexible price steady state, log markup only consists of the distortion caused by the presence of monopolistic competition. It can be shown from household's optimizing conditions that the steady-state markup is:[5]

$$\mu \equiv \log(\mathcal{M}) = w - p - mrs. \tag{9}$$

In giving an intuition behind a version of (7) without indexation ν_t, Gali (2008) notes the following: 'When the average wage in the economy is below the level consistent with maintaining (on average) the desired markup, households readjusting nominal wage will tend to increase the latter, thus generating positive wage inflation'. A similar intuition lies behind (7). We first note that average wage inflation exclusive of indexation ($\pi_t^w - \nu_t$) is identical to wage inflation as defined by Gali (2008). Thus, the intuition behind the expression (7) is as follows: when the perceived markup gap is bigger, wage setting household members see less incentive to increase nominal wages, thus resulting in less wage inflation.

[5]See Appendix B.1 for a detailed derivation.

Unemployment is introduced into the model in a way identical to that by Gali (2011). Let $l_t(i)$ be the log labour supply of individual i in the absence of real and nominal distortions. The expression for the log of individual labour supply in this case is given by the following first order condition:

$$w_t - p_t = c_t + \varphi l_t + \log(\chi_t), \tag{10}$$

where $l_t = \int_0^1 l_t(i)di$. It should be noted once again that the presence of risk sharing among individuals in a household implies that the marginal utility of consumption is equal across all individuals, further implying that $c_t = c_t(i)$. We note that the unemployment associated with labour supply $l_t(i)$ is voluntary unemployment. Also, $n_t \equiv \log(N_t)$ is the effective log labour demand under monopolistic wage setting. Using these two observations, we can define the unemployment rate as follows:

$$u_t = l_t - n_t. \tag{11}$$

Substitute (10), and (11) into the definition of average wage markup in (8) to obtain the following:

$$\mu_t = \varphi u_t. \tag{12}$$

It follows from (12) that the natural rate of unemployment is defined as follows: $u^n = (1/\varphi)\mu$. In other words, the natural rate of unemployment in a flexible price equilibrium is solely a function of wage markup.

Finally a substitution of the expressions for unemployment and its natural rate into (7) permits us to derive the NKWPC as follows:

$$\pi_t^w - \nu_t = \beta E_t(\pi_{t+1}^w - \nu_{t+1}) - \lambda\varphi(u_t - u^n). \tag{13}$$

In order to derive a reduced form version of the expression in (13) it is assumed that the unemployment gap follows the following autoregressive process of order 2 (AR(2)):[6]

$$\hat{u}_t = u_t - u^n = \phi_1 \hat{u}_{t-1} + \phi_2 \hat{u}_{t-2} + \eta_t.$$

Following Gali (2011), we suggest this process for unemployment because it seems to describe the US data quite well. Substituting this AR representation for unemployment into (13) and solving the resulting difference equation after assuming rational expectations yields the following time-varying New Keynesian Wage Phillips Curve expression:

$$\pi_t^w = \alpha_t' + \gamma_t \bar{\pi}_{t-1}^p + \phi \pi_t^z + \psi_0 u_t + \psi_1 u_{t-1} + \xi_t, \tag{14a}$$

where

$$\alpha_t' \equiv (1-\gamma_t)\pi^p + (1-\phi)\pi^z - (\psi_0 + \psi_1)u^n$$
$$\psi_0 \equiv -\frac{\lambda\varphi}{1-\beta(\phi_1+\beta\phi_2)}$$
$$\psi_1 \equiv -\frac{\lambda\varphi\beta\phi_2}{1-\beta(\phi_1+\beta\phi_2)}.$$

[6]It may be argued that this AR(2) process for unemployment seems rather ad hoc. Nevertheless, this study adopts this process in order to facilitate comparison with Gali (2011). Table 7 gives the results of the estimated process.

The random variable ξ_t is assumed to be measurement error [7] which is uncorrelated to all the other independent variables and could possibly be an autocorrelated process. It is worth noting that the sum $\gamma_t \bar{\pi}^p_{t-1} + \phi \pi^z_t$ and the function of the time-varying parameter, α'_t, are included in TV-NKWPC due to the presence of the indexed part of wages $\nu_t = x_{t|t-1}$.

We note that α_t and γ_t are negatively correlated. This property will later prove important in supporting our claim for the time-varying nature of the degree of wage indexation. The expression (14a) is a more general version of the NKWPC in that it also takes into account the time variation in the degree of wage indexation. Estimating the dynamics of wage inflation has the advantage of combining the microfounded nature of the model by Gali (2011) with the additional benefit of estimating the time variation in wage indexation. The TV-NKWPC nests the specification employed in Gali (2011) and Muto and Shintani (2014) as a special case in which the degree of wage indexation γ is assumed constant and there is no indexation to productivity (i.e. $\gamma_t = \gamma$ and $\phi = 0$). In this case, the specific form that (14a) assumes is the following expression:

$$\pi^w_t = \alpha' + \gamma \bar{\pi}^p_{t-1} + \psi_0 u_t + \psi_1 u_{t-1} + \xi_t, \tag{14b}$$

where $\alpha' = (1-\gamma)\pi^p - (\psi_0 + \psi_1)u^n$.

3 Estimating the TV-NKWPC

The empirical fit of the TV-NKWPC is investigated in this section. First, baseline estimations of the standard NKWPC are performed using US data. Diagnostic tests are then conducted on these estimations in order to look for possible evidences supporting the instability of the estimated constant term and the coefficient of inflation(α' and γ). The basic idea behind the tests is as follows: the estimates of α' and γ will not exhibit any instability if indeed the degree of wage indexation is constant. The diagnostic tests conducted can therefore be seen as indirect tests as to whether there is time variation in the degree of wage indexation. The final part of this section demonstrates the empirical fit of TV-NKWPC when estimates are conducted using US data. Not only does the use of US data facilitate comparison of the two methodologies (Gali (2011) and our study), but also it permits one to easily compare the time-varying degree of wage indexation obtained from the TV-NKWPC estimation to corresponding figures suggested by the extent of COLA coverage.

3.1 Data and preliminary evidence

This study uses quarterly data spanning the period from 1948Q1 to 2012Q4 obtained from the Bureau of Labour Statistics (BLS).[8] For the measure of inflation, Consumer Price Index (CPI) inflation is used. Wages are measured using compensation data.[9] We make use of the compensation based measure of wages in order to take advantage of its relatively longer time span. Also, according to Gali (2011), both measures yield remarkably similar results. The index of output per hour is used as a proxy for labour productivity.

[7] It has been suggested by Gali (2011) that the error term could also capture the time variation in the desired wage mark-up.

[8] Gali (2011) uses unemployment data obtained from the Haver Database.

[9] Gali (2011) makes use of earnings data in the main part of the study due to the possibility of the presence of measurement errors in compensation data.

Table 1 presents the results of the estimation. The first two columns of the baseline estimation represent a model in which wages are indexed to lagged inflation ($\bar{\pi}^p_{t-1} = \pi_{t-1}$) and a model in which wages are indexed to a moving average of lags of inflation ($\bar{\pi}^p_{t-1} = \pi^{(4)}_{t-1} = \left(\sum_{k=1}^{4} \pi_{t-k}\right)/4$). A preliminary diagnostic test run on the residuals suggests that including productivity growth adds some explanatory power to the baseline equation.[10] The regressions in the last two columns therefore include productivity growth in the list of regressors. From a theoretical point of view, it is possible that productivity enters the model by means of wage indexation, i.e. wages are indexed to lags of inflation and current productivity. It can be seen from the values of the R^2 that the fit of the model is improved when productivity growth is introduced into the model. Also, the residuals from estimations in the cases of all models shown in Table 1 display a significant level (1%) of autocorrelation.[11]

Ascari et al. (2011) document how wage indexation rises when trend inflation increases and falls when trend inflation decreases. This observation suggests the existence of instability in the NKWPC when wages are indexed to inflation. Guided by this observation, we conduct further diagnostic tests on the residuals from the regressions in Table 1 by including a nonlinear term, namely: the product of trend inflation and the measure of inflation indexed to, i.e. $\pi^{\tau}_{t-1}\pi_{t-1}$ or $\pi^{\tau}_{t-1}\pi^{4}_{t-1}$. Trend inflation is obtained by means of applying the HP filter to the quarterly inflation series. Results from Table (2) indicate a strong effect of a nonlinear term in both cases of wage indexation considered. We interpret this finding as evidence in support of our claim concerning the improved fit of the TV-NKWPC.

Finally a rolling window regression on (14a) is performed in order to obtain an idea of the time-varying parameters α_t and γ_t. This is done using the following procedure. First, the constant parameters in the expression contained in (14b) are estimated. A rolling regression is subsequently performed in order to obtain rough estimates on the parameters in the following expression:

$$x_t = \alpha_t + \gamma_t \bar{\pi}^p_{t-1} + \epsilon_t,$$

where $x_t \equiv \pi^w_t - \hat{\psi}_0 \hat{u}_t - \hat{\psi}_1 \hat{u}_{t-1}$ and ϵ_t is an independent and identically distributed (iid) zero mean normally random distributed error term. The results under the two assumptions regarding wage indexation considered are respectively presented in Figure 1 and Figure 2. In both figures, the estimated time-varying degree of wage indexation first rises to a point, after which it falls. Again, it is interesting to note that the later periods' values of wage indexation do not significantly differ from zero. This is generally in line with empirical evidence that the degree of wage indexation initially rose to high levels during the 1970s and diminished thereafter. Also, the time-varying wage indexation parameter varies between 0 and 1 in both cases. It is worth noting that the correlation between $\hat{\alpha}_t$ and $\hat{\gamma}_t$ estimated under this rolling regression technique is negative (see Table 8 in the appendix). This is expected if one holds the assumption that the reduced form specification in (14a) describes the dynamics between output and unemployment.

[10] Formal causality tests indicate that productivity growth Granger causes unemployment

[11] This is reported in Table 9 in the appendix

Figure 1: Rolling regression estimates for $\bar{\pi}_t^p = \pi_{t-1}$

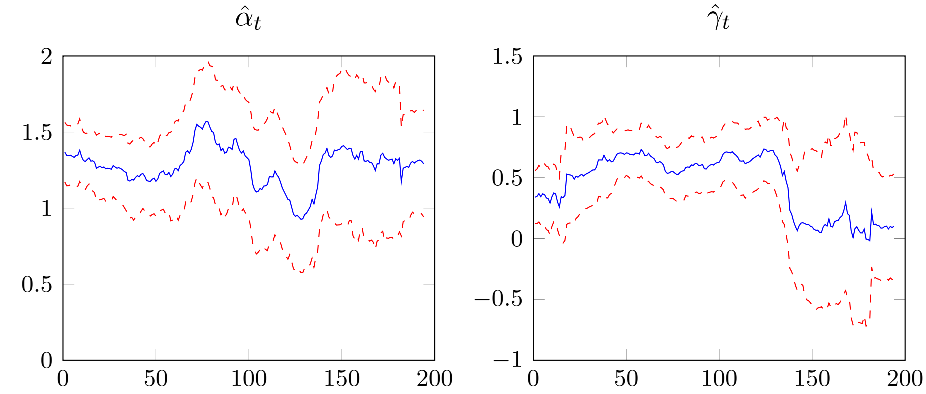


Figure 2: Rolling regressions estimates for $\bar{\pi}_t^p = \pi_{t-1}^{(4)}$

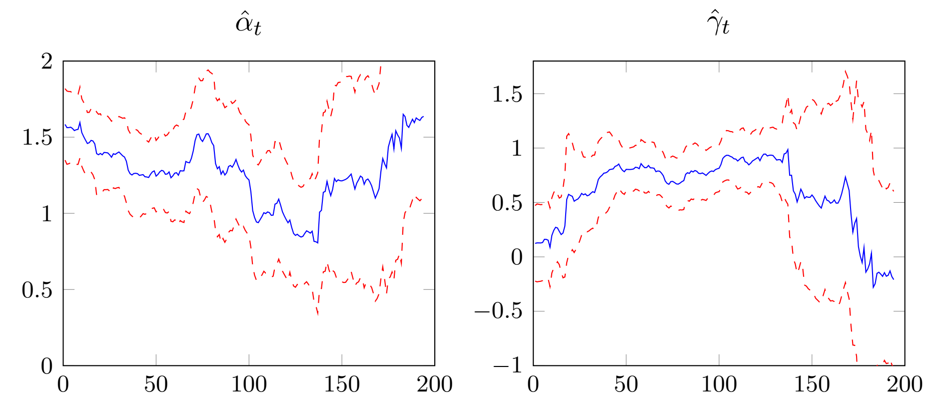


Table 1: Estimated NKWPC (π_t^w)

	(1)	(2)	(3)	(4)
u_t	-0.17	-0.33**	-0.114	-0.28**
	(0.091)	(0.094)	(0.091)	(0.093)
u_{t-1}	0.1	0.23*	0.032	0.175
	(0.092)	(0.093)	(0.091)	(0.093)
π_{t-1}	0.544**		0.582**	
	(0.050)		(0.050)	
π_{t-1}^4		0.673**		0.707**
		(0.062)		(0.061)
π_t^z			0.167**	0.15**
			(0.045)	(0.045)
const	1.217**	1.262**	1.139**	1.197**
	(0.143)	(0.143)	(0.141)	(0.142)
Adj-R^2	0.319	0.326	0.352	0.352
AIC	1.865	1.86	1.820	1.825

Estimation of $\pi_t^w = \alpha_0 + \gamma\bar{\pi}_{t-1}^p + \psi_0 u_t + \psi_1 u_{t-1} + +\psi_2 \pi_t^z + \epsilon_t$.

Standard errors of estimates are indicated in parenthesis.

* $p < 0.05$, ** $p < 0.01$.

Table 2: Estimating NKWPC residuals (ϵ_t)

	(1)	(2)	(3)	(4)
$\pi_{t-1}^\tau \pi_{t-1}$	0.323**		0.34**	
	(0.056)		(0.054)	
$\pi_{t-1}^\tau \pi_{t-1}^{(4)}$		0.420**		0.432**
		(0.074)		(0.071)
π_{t-1}	-0.518**		-0.541**	
	(0.098)		(0.095)	
$\pi_{t-1}^{(4)}$		-0.793**		-0.814**
		(0.147)		(0.143)
ϵ_{t-1}	0.111	0.117	0.074	0.099
	(0.061)	(0.061)	(0.060)	(0.061)
const	0.0120*	0.257**	0.123*	0.264**
	(0.056)	(0.72)	(0.055)	(0.070)
Adj-R^2	0.136	0.147	0.143	0.152

Estimation of $\hat{\epsilon}_t = \beta_0 + \beta_1\hat{\epsilon}_{t-1} + \beta_2\pi_{t-1}^{(p)p} + \beta_3\pi_{t-1}^\tau\pi_{t-1}^{(p)} + \nu_t$.

Standard errors of estimates are indicated in parenthesis.

* $p < 0.05$, ** $p < 0.01$.

3.2 Estimation results

In order to estimate the expression contained in (14a), we propose a state-space methodology with the time-varying degree of wage indexation (γ_t) and the measurement error (ξ_t) as the unobserved state variables. This estimation method requires one to give the law of motion for the time-varying wage indexation. As noted earlier, empirical findings suggest that wage indexation is positively correlated to trend inflation. If one assumes a simple linear relationship between wage indexation and trend inflation, it is possible to propose a highly persistent process for the wage indexation parameter.[12] It is therefore assumed that wage indexation behaves as if it were a random walk process over the sample period in consideration. Given that no restrictions are placed *a priori* with regards to the autocorrelation structure of the random process ξ_t, a stationary AR(1) process is assumed for this variable. We estimate the following empirical model:

$$\begin{cases} \pi_t^w &= \varphi_1 u_t + \varphi_2 u_{t-1} + \varphi_3 \pi_t^z + \mu_t + \varphi_4 \gamma_t + \gamma_t \bar{\pi}_{t-1}^p \\ \mu_t &= (1-\rho_\xi)\varphi_5 + \rho_\xi \mu_{t-1} + \varepsilon_t \\ \gamma_t &= \gamma_{t-1} + \eta_t \end{cases} . \tag{15}$$

where $\mu_t = \varphi_5 + \xi_t$, $\varepsilon_t \sim \ iid \ N(0, \sigma_\varepsilon^2)$ and $\eta_t \sim \ iid \ N(0, \sigma_\eta^2)$. A definition of all the coefficients contained in expression (15) above in terms of the structural parameters in the previous section is given in Table 3 below:

Table 3: Definition of coefficients

coef	definition	coef	definition
φ_1	$-\frac{\lambda\varphi}{1-\beta(\phi_1+\beta\phi_2)}$	φ_4	$-\pi^p$
φ_2	$\beta\phi_2\varphi_1$	φ_5	$(1-\phi)\pi^z + \pi^p - (\varphi_1+\varphi_2)u^n$
φ_3	ϕ		

Following Gali (2011), the TV-NKWPC is first estimated under two assumptions with regards to price inflation: that wages are indexed to a quarter lag of price inflation (π_{t-1}) and that wages are indexed to an average inflation over the last four quarters ($\pi_{t-1}^{(4)}$). Additional estimates of the TV-NKWPC are then made under the assumption that μ_t (or ξ_t) is autocorrelated, and subsequently under the assumption that μ_t (or ξ_t) is iid normal distributed. Finally, the TV-NKWPC is estimated under the assumption that there is no autocorrelation in ξ_t ($\rho_\xi = 0$) and the coefficient of lag of unemployment is zero ($\varphi_2 = 0$).

The estimations of all versions of Equation 15 were performed using 7th edition of the EVIEWS statistical package. The same package was used for all other estimations in this study except for the rolling regressions which were done in MATLAB. The results from the six estimations are presented in Table 4. There are some observations worth noting concerning the estimates of the various versions of TV-NKWPC. First, results obtained from the estimations of the various versions of the TV-NKWPC show rather striking similarities to

[12]Formal unit-root tests run on trend inflation do not statistically reject the existence of a unit-root. We acknowledge that this specification may come off as economically implausible. An alternative specification might suggest a highly persistent but stationary process (e.g. with AR(1) coefficient 0.99). We nevertheless stick to the random walk assumptions due to the following reasons: First, there is very little difference in results between a random walk specification and the persistent AR(1) specification. Secondly, it is common practice in recent literature to assume a random walk process for trend inflation.

Table 4: Estimated TV-NKWPC (π_t^w)

	$\rho_\xi \neq 0$		$\rho_\xi = 0$		$\rho_\xi = 0$, $\varphi_2 = 0$	
	π_{t-1}	$\pi_{t-1}^{(4)}$	π_{t-1}	$\pi_{t-1}^{(4)}$	π_{t-1}	$\pi_{t-1}^{(4)}$
	(1)	(2)	(3)	(4)	(5)	(6)
u_t	-0.194	-0.302**	-0.194	-0.309**	-0.114**	-0.158**
	(0.107)	(0.097)	(0.106)	(0.090)	(0.037)	(0.0249)
u_{t-1}	0.085	0.151	0.085	0.157		
	(0.106)	(0.100)	(0.105)	(0.094)		
π_t^z	0.157**	0.139**	0.158**	0.142**	0.166**	0.157**
	(0.038)	(0.038)	(0.038)	(0.387)	(0.037)	(0.039)
γ_t	0.484	-0.365*	0.434	-0.365*	0.539	-0.304**
	(0.625)	(0.145)	(0.585)	(0.128)	(0.596)	(0.108)
φ_5	1.265**	1.739**	1.282**	1.744**	1.269**	1.763**
	(0.302)	(0.167)	(0.290)	(1.156)	(0.299)	(0.155)
ρ_ξ	0	0.062				
	(0.054)	(0.051)				
$\ln(\sigma_\varepsilon^2)$	-1.264**	-1.235**	-1.264**	-1.236**	-1.261**	-1.226**
	(0.072)	(0.072)	(0.072)	(0.072)	(0.072)	(0.071)
$\ln(\sigma_\eta^2)$	-7.016**	-5.862**	-6.972	-5.840**	-7.079**	-5.884**
	(1.06)	(0.857)	(1.031)	(0.821)	(0.962)	(0.773)
AIC	1.782	1.787	1.774	1.782	1.770	1.786

[1] Estimation of the various versions of the TV-NKWPC in Equation (15).

[2] Standard errors of estimates are indicated in parenthesis

[3] * $p > 0.05$, **$p > 0.01$

those obtained from estimations of the NKWPC in Table 1. In most cases, the values of the constant terms (φ_5 in (15)) imply that the coefficients of unemployment and the coefficients of productivity are roughly similar under the various specifications.[13] An implication of these similarities could be that the error term ξ_t in the NKWPC posited to be measurement error in wage inflation by Gali (2011) is most likely explained by variations in the trend inflation (as can be seen from Table 2). Estimates for ξ_t under the various TV-NKWPC models are independent of the time-varying wage indexation and are not autocorrelated (see Figure 4). Also, estimates for the linear effect of the time-varying degree of wage indexation (φ_4 or the coefficient of γ in the table) are either not statistically significant or significantly negative as predicted by the expression (14a). [14] Finally, as indicated by the AIC values, all the versions of TV-NKWPC estimated in the table above outperform the estimation of all the versions of the NKWPC contained in Table 1. One can interpret these observations as evidence in support of the relatively better empirical fit of the TV-NKWPC to US data.

[13] This similarity only holds to the extent that wages are indexed similarly under the various specifications.

[14] The identity of the coefficient φ_4 as contained in Table 3 implies that $\varphi_4 \leq 0$ for $\pi^p \geq 0$

Furthermore, after taking into account the time variation in the degree of wage indexation when estimating the TV-NKWPC, the lag of unemployment plays no significant role in explaining wage inflation under all the versions of the TV-NKWPC estimated. This is possibly due to the fact that the persistence in wage inflation is mostly accounted for by changes in the degree of wage indexation(which is in itself a persistent process).[15] This result and the fact that ξ_t is not an autocorrelated process ($\rho_\xi = 0$ is not rejected at 10% significance level) imply that the TV-NKWPC models (5) and (6) should be preferred to the others.

The estimates for φ_4 under the aforementioned two versions of the TV-NKWPC imply two different values for non-varying steady-state inflation. In model (5), φ_4 (i.e. the coefficient of γ_t) is not statistically significant. This implies that after taking into account the effect of time-varying degree of wage indexation, the constant steady-state value of inflation is not statistically different from 0. On the other hand, the estimate of φ_4 in model (5) implies that the constant steady-state value of inflation is 0.304. When one considers (as will be shown later) that the degree of wage indexation is a function of trend inflation, it is easy to see why $\varphi_4 = 0$ is more plausible. In other words, it makes sense that the constant steady-state value of detrended inflation should be 0. Also, comparing the AIC values of models (5) and (6) suggests that one should opt for the former. Finally, the estimated time-varying degrees of wage indexation obtained under the former version of the TV-NKWPC (Figure 3) are more comparable to those suggested by COLA coverage figures. Given the result just mentioned, the next section of this study only estimates the model (5) version of the TV-NKWPC for various countries.

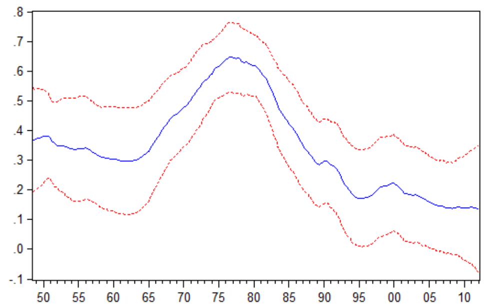
3.3 Time-varying degree of wage indexation

If the dynamics of wage inflation are indeed described by the reduced form equation (15), one would expect the state variable γ_t to effectively capture the time-varying degree of wage indexation. The estimated log variance of the γ_t's disturbance term ($\ln(\sigma_\eta^2)$) is significant at 1% under all estimated versions of TV-NKWPC. This can be interpreted as evidence in support of the time-varying nature of the degree of wage indexation. Figure 3 gives the values of the time-varying degrees of wage indexation as indicated by the smoothed estimates for γ_t under the models (5) and (6) in Table 4. The two sets of estimates for γ_t reveal a general story: the degree of wage indexation rose during the period of the Great Inflation and fell during the period of the Great Moderation, a story consistent with other empirical investigations. One main difference however exists between the two models. The magnitudes of the estimates for γ_t under model (6) slightly exceed those suggested by the proportion of workers under COLA[16] contracts. This suggests that model (5) better describes the dynamics of wage inflation.

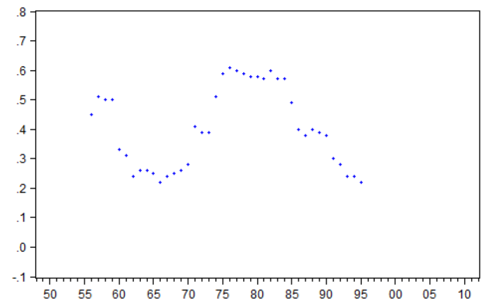
[15]This is similar to the findings of Cogley and Sbordone (2008) who argue that taking into account the variation of trend inflation makes the NKPC purely forward looking, with no need for an ad hoc backwards price indexation

[16]The COLA coverage figures are obtained from the Bureau of Labour Statistics (BLS) and Weiner (1996).

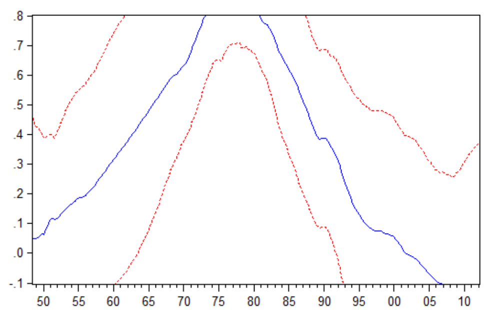
Figure 3: Smoothed estimates for γ_t



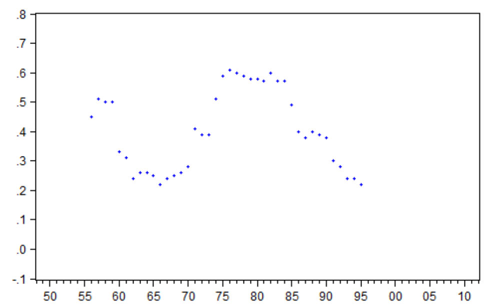
(a) Model (5)



(b) COLA coverage

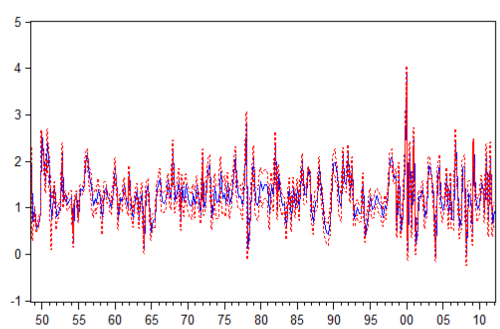


(c) Model (6)

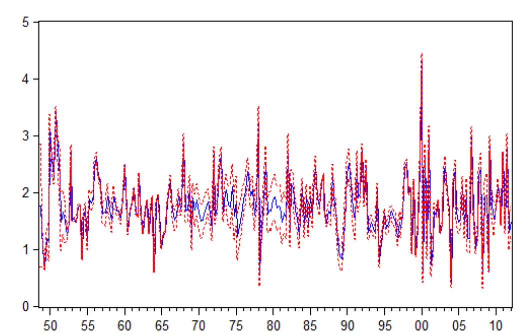


(d) COLA coverage

Figure 4: Smoothed estimates for μ_t



(a) Model (5)

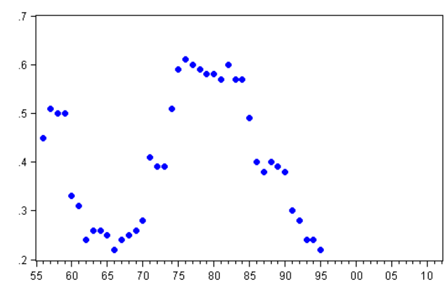


(b) Model (6)

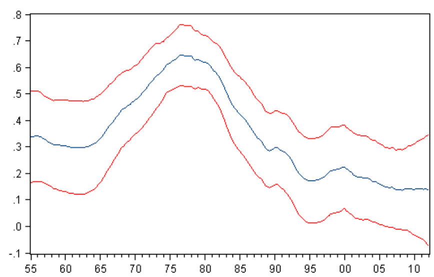
Our study is not the first attempt at estimating the time-varying degree of wage indexation. A comparison with other approaches found in existing literature reveals that our estimates for the time-varying degree of wage indexation are the closest to the figures suggested by the percentage of COLA coverage. Hofmann et al. (2010) and Ascari et al. (2011) provide estimates for the time-varying degree of wage indexation. While the estimates from their approaches produce reasonable measures for the time-varying degree of wage indexation, our approach is relatively simple, but nonetheless effectively measures this variable. Estimates for time-varying degree of wage indexation obtained in the two works just cited are compared to estimates obtained under TV-NKWPC and COLA coverage figures in Figure 5. It can be seen from this figure that the approach that best reproduces the estimates for the degree of wage indexation (γ_t) as suggested by COLA contracts coverage is the TV-NKWPC estimation. Similar to the figures suggested by COLA coverage, estimates for γ_t under the TV-NKWPC peaked at over 60% during the late 1970s and decreased to around 20% afterward. Thus, the subsequent part of this work will focus on estimating the model (5) of TV-NKWPC for selected OECD countries.

To recap, the analysis performed in this section indicates that there is indeed an empirical support for instability of the NKWPC. This instability stems from the time-varying nature of the degree of wage indexation. In particular, estimates for the time-varying degree of wage indexation (obtained from estimating the TV-NKWPC derived in the previous section) yield results strikingly similar to the percentage of COLA coverage. The latter variable is generally accepted as the proxy for the degree of wage indexation regarding US data. The estimates for the coefficients of productivity growth and unemployment under the NKWPC in Table (1) and under the TV-NKWPC in Table (4) are similar.

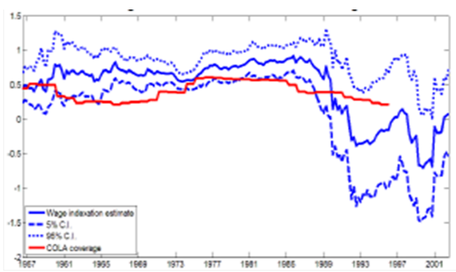
Figure 5: Various estimates for γ_t



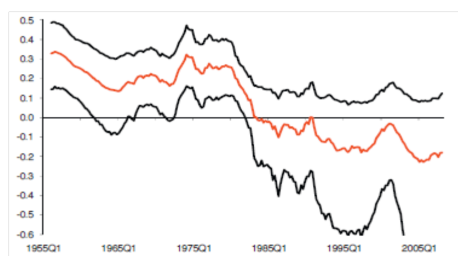
(a) COLA Coverage(BLS and Weiner (1996))



(b) TV-NKWPC



(c) Ascari et al 2011



(d) Hofmann et al, 2010

4 The TV-NKWPC in selected OECD countries

This section estimates the TV-NKWPC for 10 OECD countries and subsequently investigates the possible reasons for the time variations in the degrees of wage indexation. The countries are Austria, Belgium, Canada, Finland, Germany, Japan, Netherlands, Norway, Sweden and the United Kingdom. These countries are selected based solely on the availability of relevant data spanning a relatively long time period. For the sake of comparability, the analysis period is restricted to the period between 1970 and 2011. This is done because the data pertaining to some countries only begins from 1970.

The main variables used for the estimations performed are in most ways identical to those used in the previous section. For instance, inflation is measured by the quarter-on-quarter change in the log of CPI, while union density and union strike variables (when available) are included in the analysis as proxies for bargaining power.

However, there are some minor differences. First, hourly earnings in the manufacturing sector rather than compensation based data is used as proxy variable to measure wages. There is a possibility that this variable might not actually reflect wages in an economy dominated by the service sector. However this is the best option available as data on other potential proxy variables is scant, or in some cases, non-existent for most of the countries. Secondly, one of the following three types of unemployment data was used for the case of each country: the unemployment rate of the labour force over 15 years old, the registered unemployment rate, and the harmonized unemployment rate. Our choice of the particular type of unemployment variable for each country is motivated by the duration of the data available. It is not expected that this will qualitatively affect our result as all types of unemployment are highly correlated. Furthermore, the use of country specific OLS estimation does not require a consistent measurement of unemployment rate across countries, as a panel regression estimation would for instance. Data on unemployment rate and wages are obtained from the OECD Main Economic Indicators database. Finally, quarterly data for GDP per hour is used as the proxy for labour productivity. Data for this variable is obtainable in annual frequency from the economic data published on the website of St Louis Federal Reserve Bank. A spline interpolation is used to obtain quarterly data from available annual data.

4.1 Impressions from data

The original Phillips curve relation posits a negative relationship between wage inflation and unemployment rate. The unstable nature of the Phillips curve has often been noted by authors. Gali (2011) for instance documents this instability, especially during the period from 1970 to 1985. As a result, the correlation between wage inflation and unemployment becomes weaker when the sample period is extended to cover the period from the 1960s to the 2010s.

In order to get a crude test of the stability of the negative correlation between unemployment and wage inflation, we plot scatter diagrams depicting the relationships for each of the 10 countries. The plots in Figure 9 reveal that in most of the countries, there is at least a reasonable amount of correlation between wage inflation and unemployment. From the figure, the magnitude of the correlations between the two aforementioned variables are generally higher than 0.5. The exceptions are in the cases of Netherlands, Canada and the UK in which relatively low correlations are reported. The lowest two correlations occur in Canada and the UK. This observation could potentially hint at the poor empirical fit of the TV-NKWPC to the data of these two countries.

4.2 Results

This section presents the results obtained from the country specific estimations of the TV-NKWPC. The result for the US is included to facilitate comparison. The specific version of the TV-NKWPC estimated is repeated below:

$$\begin{cases} \pi_t^w &= \varphi_1 u_t + \varphi_3 \pi_t^z + \mu_t + \varphi_4 \gamma_t + \gamma_t \bar{\pi}_{t-1}^p \\ \mu_t &= \varphi_5 + \varepsilon_t \\ \gamma_t &= \gamma_{t-1} + \eta_t. \end{cases} \tag{16}$$

The version of the TV-NKWPC in (16) above implies that after taking into account the persistence in wage inflation accounted for by time-varying wage indexation, the possible effects of lagged unemployment are negligible. The TV-NKWPC estimated in order to investigate the robustness of our estimation excludes lagged unemployment as an explanatory variable given its low explanatory power. Table 5 gives the estimated coefficients of the TV-NKWPC for the OECD countries considered in this study.

With the exception of the UK, the estimates for the coefficients of unemployment (u_t) are significant at 1% or 5% in all countries. The magnitudes of these estimates are lowest for Sweden, Finland, the US and the UK. This may suggest the presence of a relatively higher degree of nominal wage rigidity in these countries than the others in this study. This finding is partially corroborated by Dickens et al. (2007) who find that the degree of nominal wage rigidity is indeed higher in Sweden, Finland, and the US. Also, estimates for Austria, Japan and Norway suggest that unemployment in these countries are relatively less responsive to changes in wage inflation than in the others.

There is generally no conclusive evidence in support of the explanatory role of productivity growth in the TV-NKWPC from the estimation results. For Finland, including this variable resulted in estimates for the TV-NKWPC which are difficult to explain, hence the removal of productivity growth from the list of regressors. With the exception of Belgium, Germany, Norway and the US, country specific estimates for the coefficients of productivity growth (π_t^z) imply that wages are generally more indexed to inflation than productivity growth. This still holds even when available proxies for productivity other than real GDP per hour are used. Remarkably, all the country specific estimates for the variance of the shock to the wage indexation process ($\ln(\sigma_\eta^2)$) are statistically significant at 1% . This result gives credence to the assertion that the degree of wage indexation is indeed time-varying. Furthermore, the time-varying wage indexation expression given in (4) requires the following condition to hold for the coefficient φ_4 in the presence of positive steady-state inflation: $\varphi_4 \leq 0$. This condition is due to the following identity: $\varphi_4 = -\bar{\pi}^p$. It can be seen from Table 5 that with the exception of the UK, all country specific estimates for φ_4 (the coefficient of γ_t) are either significantly negative or not statistically significant.

The constant term (φ_5) is remarkably significant and positive for all countries. In order to explain this result we recall the following definition, $\varphi_5 = (1-\phi)\pi^z + \pi^p - (\varphi_1 + \varphi_2)u^n$. In other words, φ_5 is the sum of linear functions of steady-state productivity, steady-state inflation and steady-state unemployment rate.[17] Thus, a significantly positive estimate for φ_5 in each of the countries results from the presence of positive steady-state figures for the unemployment rates and the productivity growth rates in these countries.[18]

The results indicate a good empirical fit of the TV-NKWPC to the data of the OECD countries , with the exception being the case of the UK. This is not entirely surprising as it has already been demonstrated that the correlation between wage inflation and unemployment is lowest for the UK. In spite of the poor

[17]Note that by definition $-(\varphi_1 + \varphi_2) > 0$

[18]steady-state inflation ($\bar{\pi}^p$) estimates are not statistically significant in most countries as seen from the estimates for $\varphi_4 = -\bar{\pi}^p$.

Table 5: Estimated TV-NKWPC (π_t^w)

	Estimated Coefficients						
	u_t	π_t^z	γ_t	φ_5	$\ln(\sigma_\varepsilon^2)$	$\ln(\sigma_\eta^2)$	AIC
Austria	-0.416**	0.226	-0.663**	3.888**	0.541**	-4.352**	3.619
	(0.088)	(0.21)	(0.136)	(0.477)	(0.091)	(1.135)	
Belgium	-0.274**	-0.067**	0.582	3.543**	-0.258*	-4.595**	2.938
	(0.074)	(0.159)	(0.414)	(0.717)	(0.115)	(0.642)	
Canada	-0.267*	0.2325	1.648*	2.814**	0.274 **	-5.707**	3.438
	(0.106)	(0.393)	(0.723)	(0.910)	(0.096)	(0.768)	
Finland	-0.151*		1.47	2.68*	0.803**	-5.703**	3.99
	(0.076)		(1.184)	(1.272)	(0.064)	(0.975)	
Germany	-0.214**	0.355**	-0.427**	2.912**	-0.664**	-3.292**	2.50
	(0.031)	(0.1)	(0.158)	(0.268)	(0.094)	(0.492)	
Japan	-0.659**	0.193	0.158	2.954**	0.872**	-2.433**	4.092
	(0.164)	(0.292)	(0.196)	(0.687)	(0.098)	(0.552)	
Netherlands	-0.390**	0.091	6.268	5.563**	-0.255**	-5.97**	3.19
	(0.147)	(0.233)	(4.024)	(1.429)	(0.132)	(1.075)	
Norway	-0.709*	0.768*	1.053	5.054**	0.988**	-3.344**	4.268
	(0.314)	(0.379)	(0.581)	(1.008)	(0.1)	(0.454)	
Sweden	-0.194*	0.027	2.702	3.305**	0.013	-0.168**	3.272
	(0.094)	(0.326)	(2,108)	(0.785)	(0.134)	(1.326)	
UK	-0.177	0.399	2.744*	5.368**	0.342*	-4.323**	3.77
	(0.147)	(0.363)	(1.091)	(0.944)	(0.158)	(0.578)	
the US	-0.1**	0.135*	0.765	1.213**	-1.134**	-7.187**	2.046
	(0.043)	(0.06)	(0.92)	(0.398)	(0.096)	(1.227)	

[1] The EVIEWS package used for the state-space estimation converged to two sets of estimates for the UK.The selected output shown in the table has a lower AIC value and has estimates similar to those of the model used for the robustness checks.

[2] Productivity growth was omitted from the list of regressors for Finland since including them yields unintuitive estimates for the coefficient of unemployment.

[3] Standard errors of estimates are indicated in parenthesis.

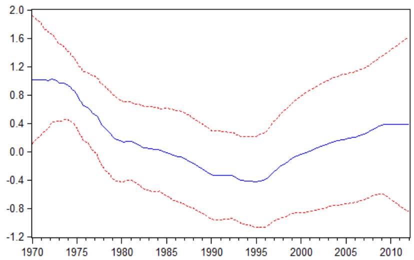
[4] * $p < 0.05$, ** $p < 0.01$

empirical fit of the TV-NKWPC to UK data, the time-varying wage indexation estimates obtained do follow the reasonable pattern of peaking in the late 1970s and falling thereafter due to fall in trend inflation since the 1970s.

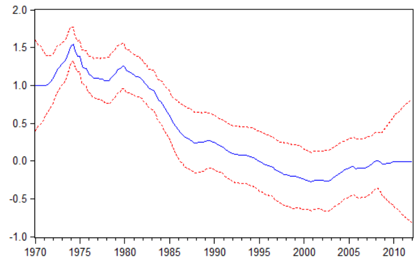
4.3 Explaining the time variation in wage indexation in OECD countries

The results obtained from estimating the TV-NKWPC as presented in Table 5 support the case for time-varying wage indexation in each of the OECD countries: the estimated log variance of the shocks to time-

Figure 6: Smoothed estimates for γ_t



(a) Austria



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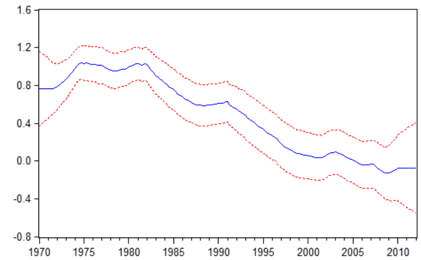

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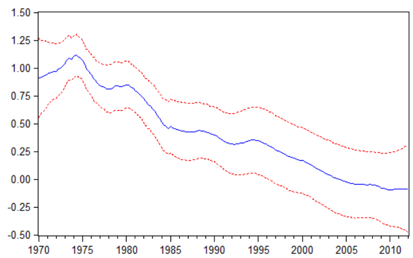
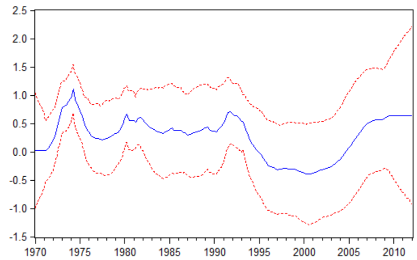


(d) Finland



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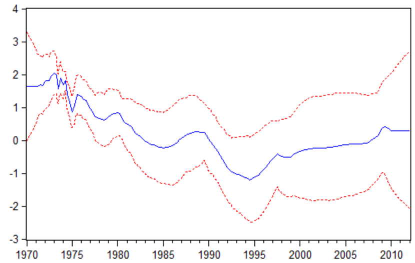

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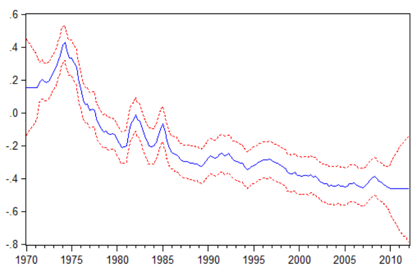
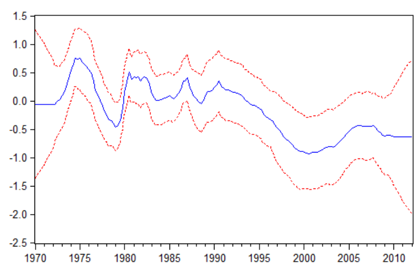
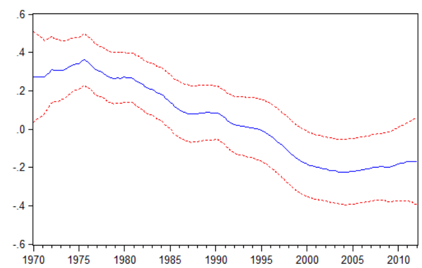


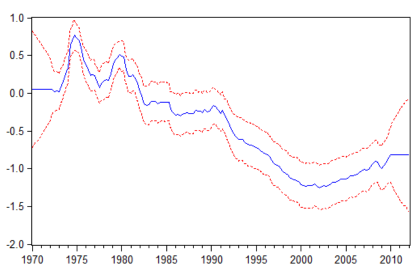
(g) Netherlands



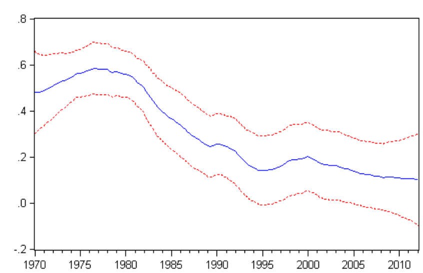
(h) Norway



(i) Sweden



(j) UK



(k) US

varying wage indexation ($\ln(\sigma_\eta^2)$) is significant at 1% level for each country. The estimated time-varying indexation ($\hat{\gamma}_t$) for each the OECD countries is presented in Figure 6. While one can be reasonably certain that the estimated degrees of wage indexation for the US do come close to the figures suggested by the generally accepted proxy for wage indexation, a similar conclusion is hard to reach for the other countries.[19]

[19]Most countries in our panel do not keep data on wage indexation. Even though data for percentage COLA coverage is available for Canada, no study has established its usefulness as a proxy for wage indexation to the best of our knowledge.

However, it can be seen from the estimations that the degree of wage indexation has been falling in the majority of these countries since the 1970s. This observation coupled with the observation that the trend inflation rates in these countries have been falling during the same period lends credence to the estimates.[20] For most of the sample period considered, the estimates for the time-varying degree of wage indexation for Austria, Germany, Japan, Norway and the UK are not statistically significant at 5%. The highest degree of wage indexation during the high inflation episode of the 1970s occurred in Belgium.[21] This result is not surprising, as this country has an automatic wage indexation policy which is applicable to all of its workers.

Having established that the estimated degrees of wage indexation (γ_t) under the TV-NKWPC do reasonably capture the degree of wage indexation, we now investigate the economic and institutional variables that explain the evolution of the degree of wage indexation. Gray (1976), Ragan and Bratsberg (2000), and Attey and de Vries (2011), among others, posit a number of variables as the factors influencing the level and distribution of the degree of wage indexation. Some of these variables are the following: real (productivity) shocks, monetary shocks (inflation uncertainty), bargaining power of unions and the number of independent unions involved in collective bargaining. In particular, the readily available estimates for time-varying wage indexation permit one to derive a test of the 'Gray hypothesis' (after Gray (1976)) which is captured in the following equation:

$$\gamma_t = f(\sigma_m^2, \sigma_z^2) \qquad \frac{\partial f}{\partial \sigma_m^2} > 0, \quad \frac{\partial f}{\partial \sigma_z^2} < 0, \tag{17}$$

where σ_m^2 denotes the variance of monetary shocks and σ_z^2 denotes the variance of real (productivity) shocks. The intuition behind this hypothesis lies in the fact that wage indexation insulates an economy from the effects of monetary shocks, while exacerbating those of real shocks. An optimal degree of indexation should therefore be close to a full indexation when monetary shocks are relatively dominant and close to zero when real shocks are relatively dominant. The aforementioned test of the Gray hypothesis can best be described as ad hoc since the original result on which the hypothesis is based describes the relationship between wage indexation under optimal monetary policy and the variances of monetary and real shocks. Therefore, any formal test of the hypothesis requires the assumption on the use of wage indexation as a policy tool in the conduct of monetary policy. However, the estimations performed in order to obtain the time-varying degrees of wage indexation variables require no such assumption. This implies that the observed time variation in wage indexation could either result from the actions of a policy maker or be an optimal outcome from bargaining between agents (for example the employers and workers unions as described in Attey and de Vries (2013)).

The wage indexation regression employs four sets of explanatory variables, namely: variances of monetary policy shocks, variances of productivity shocks, variables indicating the bargaining power of unions, and variables indicating the independence of unions. The quarter-on-quarter change in trend inflation is used

[20] Theory predicts a positive correlation between trend inflation and the degree of wage indexation.

[21] There is a general misconception that the Belgian wage legislation implies $\gamma_t = 1$ for all the time periods. However, one has to bear in mind that this full indexation represents the minimum extent of wage adjustment which cannot be undercut. Thus, it is possible to have a degree of wage indexation above 1 as observed in the late 1970s. Also, legislation put in place in 1989 imposed a maximum wage increase to be around the level of wage increase in Belgium's largest trading partners,(see Mongourdin-Denoix and Wolf (2010)). This might explain the general declining trend in wage indexation in Netherlands, Germany and France since the 1990s.

as a proxy variable for the variance monetary shocks.[22] The data for quarterly trend inflation is obtained by applying the HP filter on quarterly inflation data. In order to obtain a proxy variable for the variance of productivity shocks, a GARCH(1,1) estimation is performed on the quarterly growth in output per hour with the mean equation modeled as an AR(4) process.

The variables employed as proxies for bargaining power are the quarterly changes in union density ($\Delta unden$) and the quarterly growth rate of the number of strikes ($\Delta unstr$). Finally, to get a rough gauge of the independence of unions engaged in a wage bargaining process, we use three institutional variables namely: coordination of wage setting (crd), the predominant level at which wage bargaining occurs (lvl), and the mandatory extension of collective agreements by law to non-organized labour (ext). A high coordination of wage setting among unions, a centralized level of wage bargaining nationwide and an existence of a mandatory extension of collective agreement in one sector to other sectors generally reflect higher levels of interdependence (or lower levels of independence) among unions.

Annual data on union density and strike variables were obtained from the OECD and ILO statistics databases respectively. The annual union density data spans the period from 1960 to 2013 while the annual strike data (when available for a specific country) spans the period from early 1970s to 2013. In order to convert these to quarterly data, it is assumed that the annual data figures correspond to the last quarter of each year. We then obtain the figures for the other three quarters by the use of spline interpolation. Finally the three institutional variables used as proxies for independence of union are available from the ICTWSS database in annual frequency spanning the period from 1960 to 2011. In order to convert these variables to quarterly data, it is again assumed that the annual variables correspond to the last quarter of each year and the same figure is repeated for the previous quarters in the year. This is not only convenient but also reasonable given the fact institutional variables do not change much over time.

The country specific regression equation estimated is given below:

$$\Delta\hat{\gamma}_t = \alpha_0 + \alpha_1 \Delta\pi^{\tau}_{t-1} + \alpha_2 \sigma^2_{z,t-1} + \sum_{i=1}^{p} \beta_i barg_{i,t-1} + \sum_{j=1}^{q} \theta_j ind_{j,t} + \epsilon_{\gamma t}, \tag{18}$$

where $\epsilon_{\gamma t} \sim N(0, \sigma^2_{\gamma})$. The variables $\Delta\hat{\gamma}_t$ and σ^2_z are the quarterly changes in degree of wage indexation and quarterly variance of productivity shocks. The sets of variables denoted by $barg_i$ and ind_j represent proxies for the bargaining power of unions and independence of unions respectively. With the exception of variables used as proxies for independence of unions, all explanatory variables introduced in equation (18) are lagged. This is to account for the informational constraints faced by either policy makers or other optimizing agents (unions) when deciding the wage indexation outcome. However, these constraints do not apply to the labour market instutional variables employed in this regression due to their considerable lack of time variation.

It is expected that there is a positive relationship between trend inflation and wage indexation irrespective of whether wage indexation is derived from the conduct of optimal monetary policy or is an outcome determined by bargaining agents. This posited relationship is implied in equation (17). The aforemen-

[22] The motivation behind the use of this variable stems from the observation that higher levels of trend inflation are generally associated with higher inflation volatility (variance). Also, the use of trend inflation permits the test of whether variations in trend inflation affect the negotiations with regards to wage indexation. Finally, the use of trend inflation as a proxy for the variance of monetary policy shocks enables us to sidestep the problem of common monetary policy in Eurozone member countries.

tioned equation also predicts a negative relationship between the variance of productivity shocks and wage indexation.

In the latter case wherein wage indexation is a bargaining outcome, it is conceivable that workers index their wages to inflation in order to correct for any perceived erosion in the values of their real wages that inflation might cause. A rising trend inflation therefore increases the incidences and the extent of wage indexation. The bargaining power of unions is expected to have a positive effect on wage indexation. The independence of unions engaged in bargaining (or independence of negotiations) regarding wage indexation can have both positive and negative effects on the wage indexation outcome. For instance, the presence of mandatory extension of negotiated outcomes to all other unions might result in lower aggregate wage indexation if one bargaining process results in a lower wage indexation outcome when compared to aggregate wage indexation resulting from independent bargaining processes.

The following table gives a summary definition of the explanatory variables and the expected signs of their coefficients:

Table 6: Expected signs of coefficients

$\Delta\pi^{\tau}$	change in trend inflation	+
σ_z^2	variance of productivity shocks	-
bargaining power		
$\Delta unden$	change in union density	+
$\Delta unstri$	growth rate of union strikes	+
independence of unions		
ext	mandatory extension of settlement terms to other sectors	+/-
lvl	predominant level at which collective bargaining takes place	+/-
crd	presence of coordination in collective wage bargaining	+/-

Table (10) in the appendix contains the results of the country specific estimations of equation (18). The table indicates that in most cases, variations in the proxies for independence of unions do not explains variations in wage indexation. The only exceptions to this result are in the cases of Austria and Finland whereby coordination between negotiating unions significantly explains variations in wage indexation. The table also shows that generally, bargaining power of unions does not significantly influence the degree of wage indexation. The exceptions are the cases of Finland and Norway, in which the effects of bargaining power of unions are statistically significant in the hypothesized direction. The estimates for Belgium indicate a significant correlation between union bargaining power (as measured by union density) and wage indexation but in a direction contrary to that hypothesized.

The table provides evidence, albeit a weak one, in support of the Gray hypothesis, implying that wage indexation is decreasing in the variance of productivity shocks. The estimates have correct signs in a majority of the countries. For Austria, Canada and Finland, variance of productivity has a significant negative effect on wage indexation. In the case of Belgium and the US, however, the variance of productivity shocks have significant positive effects on wage indexation. The fact that wage indexation is automatic (given high levels of inflation resulting from stagflation or inflationary gaps) may account for the positive correlation between the lag of the variance of productivity shocks and wage indexation.

The estimated coefficient for lagged variance of productivity is significant and positive, albeit of a negligible magnitude for the US. Among all the variables consequential to explaining the time variation in wage indexation considered, variations in trend inflation is the most significant explanatory factor. The coefficients are mostly positive with the exception of those of the Netherlands and Norway. The results in the case of the Netherlands can be explained by the 'Wassenaar Agreement', which in effect moderated wages during the early 1980s when inflation was observed to be historically high. This explains the negative correlation between the lag of inflation and the degree of wage indexation.

In this section, the TV-NKWPC was estimated for 11 OECD countries. There is evidence in support of the existence of a time-varying wage indexation. The country-specific time-varying degrees of wage indexation estimated indicate the prevalence of high levels of wage indexation from the 1970s to early 1980s, and a steady decline thereafter. It can also be concluded from the estimates that variations in trend inflation significantly affect the variations in time-varying wage indexation. While there is weak evidence in support of the hypothesis that the degree of wage indexation is decreasing in the variance of productivity shocks in some countries, there is no conclusive evidence supporting the significance of labour market institutional variables in explaining wage indexation. The next section includes tests on the robustness of the results obtained to alternative specifications.

4.4 Robustness: alternative specifications to wage indexation

The relatively better empirical fit of the TV-NKWPC contained in Equation (14a) has been established by the estimations performed so far. However, this specification of the TV-NKWPC relies on the rather simple assumption of constant indexation to productivity. In order to investigate how robust the findings in the previous section are to alternative specifications, we investigate the empirical fit of the TV-NKWPC under two alternative rules for wage indexation below:

$$x_{t+k|t} = \begin{cases} 0 & k = 0 \\ \sum_{s=0}^{k-1} \left(\gamma_{t+s+1}(\bar{\pi}^p_{t+s} + \pi^z_{t+s+1}) + (1-\gamma_{t+s+1})(\pi^p + \pi^z)\right) & k \geq 1 \end{cases} \tag{19a}$$

$$x_{t+k|t} = \begin{cases} 0 & k = 0 \\ \sum_{s=0}^{k-1} \left(\gamma_{t+s+1}\bar{\pi}^p_{t+s} + (1-\gamma_{t+s+1})\pi^p + \phi_{t+s+1}\pi^z_{t+s+1} + (1-\phi_{t+s+1})\pi^z\right) & k \geq 1. \end{cases} \tag{19b}$$

The first indexation rule suggests that wages are indexed at time-varying degrees to the sum of inflation and productivity growth, while the second indexation rule implies that wages are indexed to both inflation and productivity growth at their respective time-varying degrees (γ_t and ϕ_t). The TV-NKWPC in the case of each of the wage indexation rules presented are derived in the same manner as those in the earlier sections of this paper. The reduced-form TV-NKWPC in the case of the indexation Equation (19a) is given below:

$$\pi^w_t = \alpha'_t + \gamma_t(\bar{\pi}^p_{t-1} + \pi^z_t) + \psi_0 u_t + \psi_1 u_{t-1} + \xi_t, \tag{20a}$$

where $\alpha'_t \equiv (1-\gamma_t)(\pi^p + \pi^z) - (\psi_0 + \psi_1)u^n$ and all the other parameters retain their definitions as in the TV-NKWPC expression (14a) in the main derivation. The TV-NKWPC associated with the indexation rule

(19b) is:

$$\pi_t^w = \alpha_t' + \gamma_t \bar{\pi}_{t-1}^p + \phi_t \pi_t^z + \psi_0 u_t + \psi_1 u_{t-1} + \xi_t, \tag{20b}$$

where $\alpha_t' \equiv (1-\gamma_t)\pi^p + (1-\phi_t)\pi^z - (\psi_0 + \psi_1)u^n$ and all other coefficients retain their definitions as under (14a). In expressing the TV-NKWPC contained in (20) in its state-space form , we assume that the random variable ξ_t is an iid random variable. This reflects the findings in the previous section that reject the hypothesis that ξ is an AR process. The time-varying degree of wage indexation is again assumed to behave like a random walk process in both cases of wage indexation expressions in (19). Estimating the equation (20b) requires one to specify the process for the time-varying degree of indexation of wages to productivity growth (i.e. ϕ_t). Whereas a number of studies exist that lend credence to the assertion that the degree of indexed wages to inflation (γ_t) is a function of a random walk process (for instance trend inflation), nothing in any of the available studies suggests the random walk process for the degree of wage indexation to productivity growth. Furthermore, suggesting a random walk process for ϕ_t requires one to economically justify why this variable might be non stationary. An AR(1) process with a non-zero stationary value is therefore suggested for ϕ_t. Thus, the state-space versions of 20 are given in the following two equations:

$$\begin{cases} \pi_t^w &= \varphi_1 u_t + \mu_t + \varphi_2\gamma_t + \gamma_t(\bar{\pi}_{t-1}^p + \pi_t^z) \\ \mu_t &= \varphi_4 + \xi_t \\ \gamma_t &= \gamma_{t-1} + \eta_t \end{cases} \tag{21a}$$

$$\begin{cases} \pi_t^w &= \varphi_1 u_t + \mu_t + \varphi_2\gamma_t + \varphi_3\phi_t + \gamma_t\bar{\pi}_{t-1}^p + \phi_t\pi_t^z \\ \mu_t &= \varphi_4 + \xi_t \\ \gamma_t &= \gamma_{t-1} + \eta_t \\ \phi_t &= (1-\rho_\phi)\phi + \rho_\phi\phi_{t-1} + v_t. \end{cases} \tag{21b}$$

where $\xi_t \sim N(0, \sigma_\xi^2)$, $\eta_t \sim N(0, \sigma_\eta^2)$, and $v_t \sim N(0, \sigma_v^2)$. The variable u_{t-1} is omitted among the list of regressors due to its lack of significance in explaining wage inflation as seen from Table 4. Thus, the second AR coefficient in the unemployment equation is set to zero ($\phi_2 = 0$). Also, we only consider the case where $\bar{\pi}_{t-1}^p = \pi_{t-1}$ due to the better plausibility of the estimated degree of wage indexation under this assumption compared to $\bar{\pi}_{t-1}^p = \pi_{t-1}^{(4)}$. Finally, the following definition for the other coefficients in (21) in terms of the structural parameters contained in Section 2 are given as follows:

$$\begin{aligned} \varphi_1 &= -\frac{\lambda\varphi}{1-\beta\phi_1} \\ \varphi_2 &= -\pi^p \\ \varphi_3 &= -\pi^z \\ \varphi_4 &= \pi^p + \pi^z - \varphi_1 u^n. \end{aligned}$$

The estimates of the two versions of the TV-NKWPC indicated in equation (21a) and equation (21b) are respectively presented in Table 11 and Table 12. Comparing the estimates obtained in Table 5 to those obtained in the two aforementioned tables reveal similarities of the AIC values of the three versions of the TV-NKWPC. This implies that the relative fit of the three versions of the TV-NKWPC to data are not considerably dissimilar. Furthermore, the country specific coefficients of unemployment tend to be roughly similar under the three versions of the TV-NKWPC. Finally, the figures (6), (7) and (8) do reveal generally similar trends in the variations in wage indexation.

However, it can be seen from comparing the tables (5), (11) and (12) that the AIC values given in the second of the aforementioned tables are the highest in a majority of the countries. Additionally, the magnitude of the time-varying wage estimates under the TV-NKWPC in (20a) as shown in Figure 7 differ from those under the other two models as presented in Figure 6 and Figure 8. For the US, the estimated time-varying degree of wage indexation under equation (21a) at its peak is less than half the estimates obtained under the other versions (equation (16) and equation (21b)). It should be noted that the estimates for the degree of wage indexation are closer to the figures suggested by percentage COLA coverage under the latter two models than the first model. The version of the TV-NKWPC in Equation (20a) and its estimated output will therefore be dropped from further analysis in the subsequent part of this section due to the preceding observations given in this paragraph.

Table 12 shows that incorporating a time-varying indexation to productivity growth dampens the evidence that supports the existence of time variation in wage indexation to inflation for Sweden and the UK. Moreover, the estimates in this table suggest that in the majority of the countries, the log variance of shocks to the wage indexation to productivity growth is not significant at 5% level, implying an absence of evidence for a time-varying process for indexation to productivity. The estimated coefficients of the time-varying process of wage indexation to productivity growth are statistically significant (not shown). One can therefore conclude that the specification of the TV-NKWPC captured in equation (16) does adequately describe the dynamics of wage inflation as well as the time-varying wage indexation process.

5 Conclusion and discussion

This study seeks to answer three main research questions. First, is there empirical evidence supporting the existence of time variation in wage indexation and second? Second,is there a way of estimating time-varying wage indexation using available data? Finally, what variables best explain the time variation in the degree of wage indexation? In response to the first question, this study provides ample empirical evidence to back the claim of time variation in the degree of wage indexation in 11 OECD countries. To this end, it first demonstrates the possible existence of a specification bias in the estimations carried out in Gali (2011) which are based on the assumption of a constant degree of wage indexation. A structural model incorporating time variation in the degree of wage indexation (the TV-NKWPC) is then used to estimate the degree of wage indexation. The time-varying degree of wage indexation estimates derived for the US are very similar to estimates suggested by the percentage of Cost of Living Adjustment (COLA) coverage figures, a widely accepted proxy for the time variation in wage indexation. The estimates also show a common trend of higher levels of indexation from the 1970s to early 1980s and a steady decline afterwards in the OECD countries. Furthermore, there is evidence backing the presence of 'over-indexation', i.e. when the degree of wage indexation exceeds 1, in some of the countries during the 1970s .

Subsequent analysis in the study suggests that variations in trend inflation significantly explain the variations in wage indexation in all countries. This finding is supported by Ascari et al. (2011), among others. The theoretical prediction in Attey and de Vries (2013) suggests the importance of labour market institutional variables such as independence of unions and bargaining power in explaining the level of aggregate wage indexation. However, this study yields no evidence in support of this claim in most of the countries. The estimated time-varying degree of wage indexation obtained provides us with ample opportunity to test the Gray-hypothesis (after Gray (1976)) that wage indexation is negatively correlated with the variance of productivity shocks. We uncover some evidence in support of this hypothesis. Given that no assumptions have been made concerning the derivation of wage indexation, one can interpret this result as evidence for wage indexation as possibly being the result of some optimization process which takes the stochastic structure of the economy into account.

While ours is not the first attempt to estimate the time variation in the degree of wage indexation, our estimates are more similar to the percentage COLA coverage figures than the estimates found in existing studies. The results obtained in this paper also contrasts with those obtained in Holland (1986). That paper models wage indexation as an AR(1) process.[23] The results obtained in this study imply that time variations in wage indexation are explained by variations in trend inflation. The fact that trend inflation is often empirically modeled as a random walk process supports the process proposed for wage indexation adopted in this study.

Given the empirical documentation of the time variation in wage indexation for at least the past three decades, one may wonder why the assumption of constant wage indexation seems to be the norm in macro modeling. Perhaps, the decline in trend inflation over the past two decades, and the consequent decline in the degree of wage indexation has led the attention of policy makers away from the consequences of the time variation in the degree of wage indexation. It is however still puzzling that current models that investigate the effects of rising trend inflation neglect rising levels of wage indexation since the two are often observed together. Furthermore, recent inflationary demand side policies engaged by the European Central Bank (ECB) and the FED imply that the time variation in wage indexation as a result of these policies may become of importance once again..

The TV-NKWPC model derived and estimated in this study is by no means perfect. It is possible that variations in trend inflation might not only affect the degree of wage indexation, but also how wage inflation reacts to unemployment. A possible extension of this model might adopt an approach similar to that used in Cogley and Sbordone (2008) to derive a version of NKWPC with all parameters being functions of trend inflation. With such a model, one may be able to better describe the wage dynamics in OECD countries.

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A Intratemporal decision by household members

Given the wage rate $W_t(i)$, a household member i maximizes labour income subject to the constraint implied by the aggregate labour. The Lagrangian formulation of this intratemporal problem is given as follows:

$$\max_{N_t(i)} W_t(i)N_t(i)di - \lambda\left(N_t - \left[\int_0^1 N_t(i)^{1-\frac{1}{\epsilon}}di\right]^{\frac{\epsilon}{\epsilon-1}}\right). \tag{A.1}$$

We note that λ is a constant since this is a simple intratemporal (static) optimization problem. Noting that the constraint implied by the aggregate labour is binding permits one to write the first order conditions associated with this problem in addition to other implied derivations as in the following expressions:

$$W_t(i) = \lambda N_t(i)^{-\frac{1}{\epsilon}} N_t^{\frac{1}{\epsilon}}$$
$$W_t(j) = \lambda N_t(j)^{-\frac{1}{\epsilon}} N_t^{\frac{1}{\epsilon}}$$
$$\frac{W_t(i)}{W_t(j)} = \left(\frac{N_t(i)}{N_t(j)}\right)^{-\frac{1}{\epsilon}}.$$

The final expression is derived by dividing the first expression by the second. Assuming that $N_t(j) = N_t$, then $W_t(j) = W_t$. Noting this allows one to derive the demand for individual labour type as follows:

$$N_t(i) = \left(\frac{W_t(i)}{W_t}\right)^{-\epsilon} N_t. \tag{A.2}$$

As an intermediate step, both sides of the expression (A.2) are raised to the power $\frac{\epsilon}{\epsilon-1}$. The expression for aggregate wages W_t is then derived in the following expressions:

$$N_t(i)^{\frac{\epsilon-1}{\epsilon}} = N_t^{\frac{\epsilon-1}{\epsilon}} W_t^{\epsilon-1} W_t(i)^{1-\epsilon}$$
$$\int_0^1 N_t(i)^{1-\frac{1}{\epsilon}}di = N_t^{\frac{\epsilon-1}{\epsilon}} W_t^{\epsilon-1} \int_0^1 W_t(i)^{1-\epsilon}di.$$

To proceed further, we begin by making the following substitution as implied by the aggregate labour index: $N_t^{\frac{\epsilon-1}{\epsilon}} = \int_0^1 N_t(i)^{1-\frac{1}{\epsilon}}di$. This permits us to derive the following expression for aggregate wages.

$$W_t^{1-\epsilon} = \int_0^1 W_t(i)^{1-\epsilon}di.$$

The final expression can be rearranged to give the definition for aggregate wages W_t as follows:

$$W_t = \left[\int_0^1 W_t(j)^{1-\epsilon}di\right]^{\frac{1}{1-\epsilon}}. \tag{A.3}$$

B The time-varying New Keynesian Wage Phillips Curve

B.1 Deriving the structural TV-NKWPC

The problem of a worker optimizing in the current period is to choose the optimal wage rate (W_t^*) in order to maximize their utility subject to their budget constraints and their labour demand schedules. In algebraic terms, the problem of the reoptmizing household is to maximize:

$$E_t\left[\sum_{k=0}^{\infty}(\beta\theta)^k U(C_{t+k|t}, N_{t+k|k})\right] \tag{B.1}$$

subject to the aggregate labour demand constraint and the budget constraint given respectively below:

$$N_{t+k|t} = \left(\frac{W_t^* X_{t+k|t}}{W_{t+k}}\right)^{-\epsilon} N_{t+k} \tag{B.2}$$

$$P_{t+k}C_{t+k|t} + E_{t+k}\{Q_{t+k,t+k+1}B_{t+k+1|t} \leq B_{t+k|t} + W_t^* X_{t+k|t}N_{t+k|t} - T_{t+k}. \tag{B.3}$$

Noting that $C_{t+k|t}$ and $N_{t+k|k}$ are both functions of W_t^*, one can derive the first order condition associated with this problem as follows:

$$\begin{aligned}
0 =& E_t\left[\sum_{k=0}^{\infty}(\beta\theta)^k\left(\frac{\partial U}{\partial C_{t+k|t}}\frac{\partial C_{t+k|t}}{\partial W_t^*} + \frac{\partial U}{\partial N_{t+k|t}}\frac{\partial N_{t+k|t}}{\partial W_t^*}\right)\right] \\
=& E_t\left[\sum_{k=0}^{\infty}(\beta\theta)^k\left((1-\epsilon)N_{t+k|t}U_C(\cdot)\frac{X_{t+k|t}}{P_{t+k}} - \epsilon N_{t+k|t}U_N(\cdot)\frac{1}{W_t^*}\right)\right] \\
=& E_t\left[\sum_{k=0}^{\infty}(\beta\theta)^k N_{t+k|t}U_C(\cdot)\left(W_t^*\frac{X_{t+k|t}}{P_{t+k}} - \frac{\epsilon}{1-\epsilon}\frac{U_N}{U_C}\right)\right].
\end{aligned}$$

Let the marginal rate of substitution for any household member that resets its wages in time t be defined as $MRS_{t+k|t} = -U_N/U_C$, and let $\mathcal{M} = \epsilon/(\epsilon-1)$. The last expression then becomes:

$$\sum_{k=0}^{\infty}(\beta\theta)^k E_t\left[N_{t+k|t}U_c(\cdot)\left(\frac{W_t^* X_{t+k|t}}{P_{t+k}} - \mathcal{M}MRS_{t+k|t}\right)\right] = 0. \tag{B.4}$$

There are a couple of points worth noting about the non stochastic steady state version of B.4 which will be useful for the derivation of the loglinearized version of this equation.

- While prices (P) and wages (W) may be non stationary even in the steady state, real wages (W/P) are stationary since consumption (C) and labour (N) are stationary in the non-stochastic steady state. This further implies that the marginal rate of substitution is also stationary.

- The steady-state value of the indexed part of wages is $X_{t+1|t}$ is 1. Also the definition of the steady state (absence of any form of nominal rigidity) implies that there is no indexation ($X = 1$or $x = 0$).

- The non-stochastic steady-state version of this equation implies that the following holds:

$$\frac{W}{P} = \mathcal{M}MRS$$

Let $\mu = log\mathcal{M}$. In terms of log variables the last expression can be written as follows:

$$\begin{aligned}
w - p &= \mu + mrs \\
\mu &= w - p - mrs.
\end{aligned}$$

One can then loglinearize equation (B.4) as shown in the following steps.

$$
\begin{aligned}
0 =& \sum_{k=0}^{\infty}(\beta\theta)^k E_t\left(\frac{W}{P}(w_t^* - w) + \frac{W}{P}x_{t+k|t} - \frac{W}{P}(p_{t+k} - p) - \mathcal{M}MRS(mrs_{t+k|t} - mrs)\right)\\
=& \sum_{k=0}^{\infty}(\beta\theta)^k E_t\left(\frac{W}{P}(w_t^* - w) + \frac{W}{P}x_{t+k|t} - \frac{W}{P}(p_{t+k} - p) - \frac{W}{P}(mrs_{t+k|t} - mrs)\right)\\
=& \sum_{k=0}^{\infty}(\beta\theta)^k E_t\left((w_t^* - w) + x_{t+k|t} - (p_{t+k} - p) - (mrs_{t+k|t} - mrs)\right)\\
=& \sum_{k=0}^{\infty}(\beta\theta)^k E_t\left(w_t^* + x_{t+k|t} - p_{t+k} - mrs_{t+k|t} - (w + p - mrs)\right)\\
=& \sum_{k=0}^{\infty}(\beta\theta)^k E_t\left(w_t^* + x_{t+k|t} - p_{t+k} - mrs_{t+k|t} - \mu\right)\\
=& w_t^*/(1-\beta\theta) + \sum_{k=0}^{\infty}(\beta\theta)^k E_t\left(x_{t+k|t} - p_{t+k} - mrs_{t+k|t} - \mu\right).
\end{aligned}
$$

The final expression implies the following expression for optimal wages set by members of the household who have the opportunity to set wages:

$$
w_t^* = (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^k E_t\left(mrs_{t+k|k} + p_{t+k} - x_{t+k|t} + \mu\right).
$$

The next step is to express marginal rate of substitution in terms of wages and the indexed part of wages. We begin by noting that due to perfect risk sharing by members of the household, all members have identical marginal utility hence identical consumption $C_{t+k} = C_{t+k|k}$. Let the utility function of a representative household be:

$$
U(C_t, N_t(i), \chi_t) = \log C_t - \chi_t \int_0^1 \frac{N_t(i)^{1+\varphi}}{1+\varphi} di.
$$

The derivation of the expression of marginal rate of substitution of a household member in period $t+k$ given that they last set their optimal wage rate in period t is given in the following steps:

$$
\begin{aligned}
MRS_{t+k|t} &= -U_N/U_C\\
&= \chi_{t+k} C_{t+k|t} N_{t+k|t}^{\varphi}\\
&= \chi_{t+k} C_{t+k} N_{t+k|t}^{\varphi}\\
mrs_{t+k|t} &= \log(\chi) + c_{t+k} + \varphi n_{t+k|t}\\
&= \log(\chi) + c_{t+k} + \varphi n_{t+k} + (\varphi n_{t+k|t} - \varphi n_{t+k})\\
&= mrs_{t+k} + \varphi(n_{t+k|t} - n_{t+k}).
\end{aligned}
$$

The loglinearized version of (B.2) implies $\varphi(n_{t+k|t} - n_{t+k}) = -\varphi\epsilon(w_t^* + x_{t+k|t} - w_{t+k})$. Making this substitution permits the last expression for $mrs_{t+k|t}$ to be written as follows:

$$
mrs_{t+k|t} = mrs_{t+k} - \epsilon\varphi(w_t^* + x_{t+k|t} - w_{t+k}). \tag{B.5}
$$

We proceed further by expressing optimal wages by wage setting household members as a function of aggregate marginal rate of substitution and other variables.

$$
\begin{aligned}
w_t^* =& (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^k E_t\left[mrs_{t+k|k} + p_{t+k} - x_{t+k|t} + \mu\right] \\
=& (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^k E_t\left[mrs_{t+k} - \epsilon\varphi(w_t^* + x_{t+k|t} - w_{t+k}) + p_{t+k} - x_{t+k|t} + \mu\right] \\
=& -\epsilon\varphi w_t^* + (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^k E_t\left[mrs_{t+k} - w_{t+k} + p_{t+k} + \mu + (1+\epsilon\varphi)(w_{t+k} - x_{t+k|t})\right] \\
w_t^* =& (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^k E_t\left[(mrs_{t+k} - w_{t+k} + p_{t+k} + \mu)/(1+\epsilon\varphi) + w_{t+k} - x_{t+k|t}\right].
\end{aligned}
$$

Let $w_t - p_t - mrs_t = \mu_t$ and $\mu_t - \mu = \hat{\mu}_t$. Then

$$
w_t^* = (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^k\left(w_{t+k} - x_{t+k|t} - \frac{1}{1+\epsilon\varphi}\hat{\mu}_{t+k}\right). \tag{B.6}
$$

Noting that $w_t^* = w_{t|t}$ and that $x_{t+1|t} = (x_{t+k+1|t} - x_{t+k+1|t+1})$, a step by step derivation of an intermediate version of the structural NKWPC can be given as follows:

$$
\begin{aligned}
w_{t+1|t+1} =& (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^k E_{t+1}\left(w_{t+k+1} - x_{t+k+1|t} - \frac{1}{1+\epsilon\varphi}\hat{\mu}_{t+k+1} + (x_{t+k+1|t} - x_{t+k+1|t+1})\right) \\
=& (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^k E_{t+1}\left(w_{t+k+1} - x_{t+k+1|t} - \frac{1}{1+\epsilon\varphi}\hat{\mu}_{t+k+1}\right) + (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^k\nu_{t+1} \\
=& (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^k E_{t+1}\left(w_{t+k+1} - x_{t+k+1|t} - \frac{1}{1+\epsilon\varphi}\hat{\mu}_{t+k+1}\right) + \nu_{t+1},
\end{aligned}
$$

where $\nu_{t+1} = x_{t+1|t}$. Multiplying both sides of the last expression by $\beta\theta$ and subsequently taking expectation conditional on information available at time t, we get the following:

$$
\begin{aligned}
(\beta\theta)w_{t+1}^* =& (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^{k+1} E_{t+1}\left(w_{t+k+1} - x_{t+k+1|t} - \frac{1}{1+\epsilon\varphi}\hat{\mu}_{t+k+1}\right) + (\beta\theta)\nu_{t+1} \\
(\beta\theta)E_t w_{t+1}^* =& (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^{k+1} E_t\left(w_{t+k+1} - x_{t+k+1|t} - \frac{1}{1+\epsilon\varphi}\hat{\mu}_{t+k+1}\right) + (\beta\theta)E_t\nu_{t+1}.
\end{aligned}
$$

Next, we replace the time index by making the substitution $s = k+1$. This implies the previous expression can be alternatively rendered as:

$$
(\beta\theta)E_t w_{t+1}^* = (1-\beta\theta)\sum_{s=1}^{\infty}(\beta\theta)^s E_t\left(w_{t+s} - x_{t+s|t} - \frac{1}{1+\epsilon\varphi}\hat{\mu}_{t+s}\right) + (\beta\theta)E_t\nu_{t+1}.
$$

Finally, the equation for optimal wages for wage setting household members at time t given in (B.6) implies the previous equation can be recast as follows:

$$
(\beta\theta)E_t w_{t+1}^* = w_t^* - (1-\beta\theta)\left(w_t - x_{t|t} - \frac{1}{1+\epsilon\varphi}\hat{\mu}_t\right) + (\beta\theta)E_t\nu_{t+1}.
$$

We recall from the wage indexation expression in (4) given in the main part of this work that$x_{t|t} = 0$. Noting this, the last expression can be rearranged to result in the following expression:

$$
w_t^* = \beta\theta E_t(w_{t+1}^* - \nu_{t+1}) + (1-\beta\theta)(w_t - (1+\epsilon\varphi)^{-1}\hat{\mu}_t). \tag{B.7}
$$

Aggregate wages in the economy is assumed to be a weighted average of reset wages and indexed wages (those not derived from optimizing). The expression for aggregate wages and the loglinearized version is presented below:

$$\begin{aligned}
W_t &= \left[\theta\left(W_{t-1}X_{t|t-1}\right)^{1-\epsilon} + (1-\theta)(W_t^*)^{1-\epsilon}\right]^{\frac{1}{1-\epsilon}} \\
W_t^{1-\epsilon} &= \theta(W_{t-1}X_{t|t-1})^{1-\epsilon} + (1-\theta)(W_t^*)^{1-\epsilon} \\
(1-\epsilon)W^{1-\epsilon}(w_t - w) &= (1-\epsilon)\left[W^{1-\epsilon}\theta[(w_{t-1}-w) + x_{t|t-1}] + (1-\theta)W^{1-\epsilon}(w_t^* - w)\right] \\
w_t - w &= \theta[(w_{t-1}-w) + x_{t|t-1}] + (1-\theta)(w_t^* - w)
\end{aligned}$$

which after making the substitution $\nu_t = x_{t|t-1}$ leads us to this expression

$$w_t = \theta(w_{t-1} + \nu_t) + (1-\theta)w_t^*. \tag{B.8}$$

Substitution of the above expression into that in B.7 gives the following version of the NKWPC:

$$\pi_t^w - \nu_t = \beta E_t(\pi_{t+1}^w - \nu_{t+1}) - \lambda\hat{\mu}_t \qquad \lambda = \frac{(1-\theta)(1-\beta\theta)}{(1+\epsilon\varphi)\theta}. \tag{B.9}$$

B.2 Reduced-form TV-NKWPC

Next we derive the reduced form TV-NKWPC. To do this, we introduce unemployment into equation (B.9) by noting that $\hat{\mu}_t = \varphi\hat{u}_t$ as explained in the main part of this text. We assume unemployment is an AR(2) process given as follows:

$$\hat{u}_t = \phi_1\hat{u}_{t-1} + \phi_2\hat{u}_{t-2} + v_t \qquad v_t \sim N(0, \sigma_v^2).$$

Let $V_t = \pi_t^w - \nu_t$, and $\delta = \lambda\varphi$. We rewrite the expression (B.9) as follows:

$$V_t = \beta E_t V_{t+1} - \delta\hat{u}_t.$$

To solve the difference equation we make an initial guess. We guess that V_t will be a function of unemployment and its lag. Thus,

$$V_t = \psi_0\hat{u}_t + \psi_1\hat{u}_{t-1}.$$

We lead V_t by one time period and take expectation of the resulting expression. This derives the following sets of equations:

$$\begin{aligned}
\beta E_t V_{t+1} &= \beta\psi_0 E_t(\hat{u}_{t+1}) + \beta\psi_1\hat{u}_t \\
&= \beta\psi_0(\phi_1\hat{u}_t + \phi_2\hat{u}_{t-1}) + \beta\psi_1\hat{u}_t \\
&= (\beta\psi_0\phi_1 + \beta\psi_1)\hat{u}_t + \beta\psi_0\phi_2\hat{u}_{t-1}.
\end{aligned}$$

We substitute the previous expression for $\beta E_t V_{t+1}$ into our initial guess $V_t = \beta E_t V_{t+1} - \delta\hat{u}_t$ to obtain the following:

$$V_t = (\beta\psi_0\phi_1 + \beta\psi_1 - \delta)\hat{u}_t + \beta\psi_0\phi_2\hat{u}_{t-1}.$$

Equating this expression to the initial guess $V_t = \psi_0 \hat{u}_t + \psi_1 \hat{u}_{t-1}$ results in the following simultaneous equation for the coefficients ψ_0 and ψ_1:

$$\psi_1 = (\beta \psi_0 \phi_2)$$
$$\psi_0 = (\beta \psi_0 \phi_1 + \beta \psi_1 - \delta).$$

Solving the simultaneous equations above yield the following expressions for ψ_0 and ψ_1:

$$\psi_0 = -\frac{\delta}{1 - \beta(\phi_1 + \beta \phi_2)}$$
$$\psi_1 = -\frac{\beta \phi_2 \delta}{1 - \beta(\phi_1 + \beta \phi_2)}.$$

After making the substitutions $\hat{u}_t = (u_t - u^n)$ and assuming the presence of a measurement error ξ_t, the reduced form TV-NKWPC can be written as:

$$\pi_t^w = (1 - \gamma_t)\pi^p + (1 - \phi)\pi^z + \gamma_t \bar{\pi}_{t-1}^p + \phi \pi_t^z + \psi_0(u_t - u^n) + \psi_1(u_{t-1} - u^n) + \xi_t. \tag{B.10}$$

C Tables and figures

C.1 Tables

Table 7: AR(2) process for unemployment (u_t)

parameter	const	ϕ_1	ϕ_2	Adj R^2
	0.299**	1.470**	-0.520**	0.953
std err	(0.082)	(0.054)	(0.054)	

* $p < 0.05$, ** $p < 0.01$

Table 8: Estimating $\hat{\alpha}_t = \beta_0 + \beta_1 \hat{\gamma}_t + \varepsilon_t$

	$\bar{\pi}^p_{t-1} = \pi^{(4)}_{t-1}$			$\bar{\pi}^p_{t-1} = \pi_{t-1}$		
	β_0	β_1	R^2	β_0	β_1	R^2
estimate	1.375 **	-0.217 **	0.167	1.566 **	-0.497 **	0.584
std err	0.02	0.03		0.0179	0.0351	

[1] This table estimates the correlation between the time-varying parameters $\hat{\alpha}_t$ and $\hat{\gamma}_t$ obtained from rolling regression estimates of Equation (14a).

[2] * $p < 0.05$, ** $p < 0.01$

Table 9: Autocorrelation in $\hat{\epsilon}$

	(1)	(2)	(3)	(4)
$\hat{\epsilon}_{t-1}$	0.18**	0.21**	0.14*	0.20**
	(0.062)	(0.062)	(0.062)	(0.062
const	0.002	0.003	0.002	0.003
	(0.037)	(0.036)	(0.037)	(0.036)
Adj-R^2	0.029	0.042	0.016	0.036
AIC	1.815	1.792	1.77	1.752

Estimation of $\hat{\epsilon}_t = \zeta_0 + \zeta_1 \hat{\epsilon}_{t-1} + v_t$

Standard errors of estimates are indicated in parenthesis.

* $p < 0.05$, ** $p < 0.01$.

Table 10: Determinants of wage indexation $\hat{\gamma}_t$ Equation (18)

	Estimated Coefficients								
				Union Bargaining Power		Union Independence			
	const	$\Delta\pi^{\tau}_{t-1}$	$\sigma^2_{z,t-1}$	$\Delta unden_{t-1}$	$\Delta unstr_{t-1}$	crd_t	ext_t	lvl_t	*Adj* R^2
Austria	0.146**	0.180**	-0.907**	0.005		-0.031**		-0.0003	0.48
	(0.019)	(0.0.061)	(0.249)	(0.005)		(0.004)		(0.004)	
Belgium	0.007	0.499**	0.309*	-0.017*		-0.005		0.001	0.33
	(0.017)	(0.065)	(0.144)	(0.007)		(0.004)		(0.004)	
Canada	-0.004	0.267**	-0.663**	0.004	-0.005	-0.002		0.007	0.50
	(0.004)	(0.034)	(0.170)	(0.003)	(0.005)	(0.001)		(0.003)	
Finland	0.001	0.091**	0.015	0.008**		0.004**		-0.005**	0.24
	(0.003)	(0.027)	(0.013)	(0.002)		(0.001)		(0.001)	
Germany	-0.023	1.144**	-0.081	0.02		0.01			0.24
	(0.033)	(0.167)	(0.065)	(0.013)		(0.009)			
Japan	0.085	0.347*	-0.821	-0.001	-0.259	-0.018			0.03
	(0.085)	(0.215)	(0.602)	(0.081)	(0.155)	(0.169)			
Netherlands	-0.008	-0.172**	-0.413	0.005		-0.003		0.005	0.09
	(0.014)	(0.053)	(0.219)	(0.008)		(0.003)		(0.004)	
Norway	0.117	-0.401*	-3.605**	0.09**		-0.009	-0.012	-0.006	0.245
	(0.044)	(0.155)	(0.995)	(0.017)		(0.008)	(0.013)	(0.009)	
Sweden	0.014*	0.179**	0.06	0.005		-0.004		-0.001	0.16
	(0.007)	(0.043)	(0.133)	(0.004)		(0.002)		(0.002)	
UK	0.001	0.327**	1.056	0.026	0.06	0.013		-0.013	0.22
	(0.008)	(0.1)	(0.656)	(0.013)	(0.068)	(0.006)		(0.006)	
the US	-0.005**	0.113**	0.006**			-0.0001			0.52e
	(0.001)	(0.014)	(0.002)			(0.0004)			

[1] Estimation of Equation (18).
[2] Explanations regarding the explanatory variables are found in Table 6 in the main text.
[3] $*p < 0.05$, $**p < 0.01$., Standard errors of estimates are indicated in parenthesis

Table 11: Robustness: Estimating version (21a) of the TV-NKWPC

	Estimated Coefficients					
	u_t	γ_t	φ_4	$\ln(\sigma_\varepsilon^2)$	$\ln(\sigma_\eta^2)$	AIC
Austria	-0.422**	-1.253**	4.062**	0.568**	-5.249**	3.633
	(0.098)	(0.377)	(0.502)	(0.089)	(1.257)	
Belgium	-0.313**	1.236	4.005**	-0.372*	-4.291**	3.020
	(0.100)	(0.717)	(0.807)	(0.150)	(0.379)	
Canada	-0.303**	1.683*	3.252**	0.266 **	-5.758**	3.432
	(0.088)	(0.766)	(0.829)	(0.095)	(0.654)	
Finland	-0.121*	0.504	2.463**	0.665**	-5.494**	3.806
	(0.073)	(1.127)	(0.727)	(0.059)	(0.783)	
Germany	-0.211**	-0.880**	3.085**	-0.614**	-4.549**	2.513
	(0.036)	(0.288)	(0.319)	(0.107)	(0.645)	
Japan	-0.701**	-0.345**	3.252**	0.742**	-1.772**	4.085
	(0.184)	(0.096)	(0.687)	(0.119)	(0.510)	
Netherlands	-0.3373**	3.565*	4.728**	-0.321*	-5.112**	3.145
	(0.144)	(1.700)	(0.957)	(0.141)	(0.652)	
Norway	-0.593*	0.280	4.051**	1.001**	-3.247**	4.261
	(0.302)	(0.668)	(0.958)	(0.096)	(0.564)	
Sweden	-0.187*	2.787	3.156**	0.029	-6.598**	3.261
	(0.085)	(2.362)	(0.680)	(0.137)	(1.226)	
UK	-0.106	2.33*	4.737**	0.418**	-4.527**	3.800
	(0.132)	(0.968)	(0.835)	(0.143)	(0.617)	
the US	-0.11*	9.782	0.281	-1.159**	-9.616**	2.046
	(0.043)	(8.569)	(1.359)	(0.101)	(1.662)	

[1] Standard errors of estimates are indicated in parenthesis

[2] * $p < 0.05$, ** $p < 0.01$

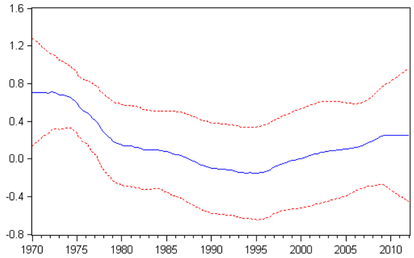
Table 12: Robustness: Estimating version (21b) of the TV-NKWPC

	Estimated Coefficients									
	u_t	γ_t	ϕ_t	φ_4	ρ_ϕ	ϕ	$\ln(\sigma^2_\varepsilon)$	$\ln(\sigma^2_\eta)$	$\ln(\sigma^2_v)$	AIC
Austria	-0.497**	1.444	0.270	4.323**	-0.231*	0.017	-0.591	-5.321**	0.003	3.561
	(0.161)	(1.503)	(0.247)	(1.075)	(0.092)	(0.252)	(0.365)	(1.320)	0.481	
Belgium	-0.256*	1.614 *	5.808	3.085**	-0.174**	0.007	25.165	-4.686**	-4.013	2.925
	(0.064)	(0.736)	(3.693)	(1.1363)	(0.061)	(0.156)	(3e9)	(0.560)	(1.156)	
Canada	-0.268*	1.683*	-8.567	4.687	0.027	0.212	0.270	(5.682**	-25.229	3.476
	(0.106)	(0.766)	(113.15)	(24.597)	(44e4)	(0.391)	(1.157)	(0.791)	1.92e9	
Finland	-0.090	1.311	5.272	1.928	-0.274	-0.061	-17.831	-5.895**	-2.984*	3.77
	(0.061)	(1.614)	(6.719)	(1.663)	(0.231)	(0.208)	(1.09e8)	(1.0756)	(1.417)	
Germany	-0.195**	-0.349** ‘	0.619	2.376**	-0.185	0.47**	-1.248**	-3.575**	-2.026	2.499
	(0.031)	(0.147)	(0.862)	(0.545)	(0.221)	(0.151)	(0.407)	(0.645)	(1.051)	
Japan	-0.607**	0.251	27.001	-2.137	-0.144*	0.179	-25.021	-3.189**	-5.747	4.096
	(0.161)	(0.247)	(44.512)	(13.466)	(0.057)	(0.258)	(2.2e9)	(0.471)	(3.213) ‘	
Netherlands	-0.358**	8.479	0.046	6.087**	0.269	0.093	-0.671**	-7.25**	-0.152	3.136
	(0.128)	(6.705)	(0.277)	(1.981)	(0.138)	(0.337)	(0.178)	(1.406)	(0.651)	
Norway	-0.789**	0.757	5.133	-0.21	-0.324	0.976**	-21.117	-2.37**	-2.804	4.23
	(0.295)	(0.412)	(6.988)	(7.244)	(0.287)	(0.323)	3e9	(0.306)	(1.52)	
Sweden	-0.173	2.709	-19.281	3.525	–0.157	0.038	-22.718	-5.959	-5.909	3.313
	(0.085)	(2.202)	(63.638)	(6.863)	(0.142)	(0.326)	(1e9)	(1.296)	(6.758)	
UK	-0.162	2.836**	-0.63**	5.384**	-0.466**	0.447	-0.178	-4.246	0.646	3.753
	(0.132)	(1.026)	(0.077)	(1.004)	(0.174)	(0.357)	(0.257)	(0.529)	(0.52)	
the US	-0.097*	0.925	-0.934**	-1.288**	0	0.15*	-1.4*	-7.43**	-2.25**	2.068
	(0.048)	(1.038)	(0.351)	(0.403)	(0.032)	(0.073)	(0.207)	(1.214)	(0.618)	

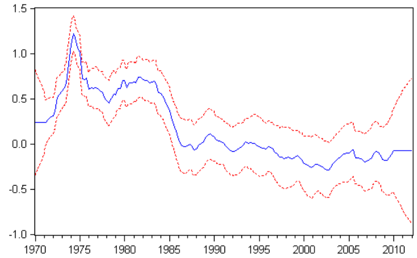
[1] Standard errors of estimates are indicated in parenthesis
[2] * $p < 0.05$, ** $p < 0.01$

C.2 Figures

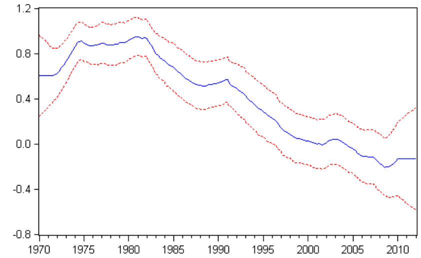
Figure 7: Smoothed estimates for γ_t: Model (21a)



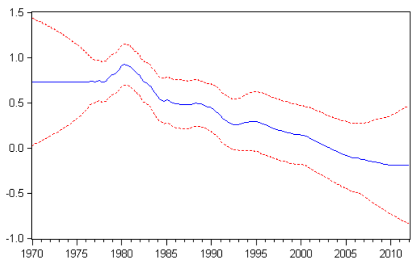
(a) Austria



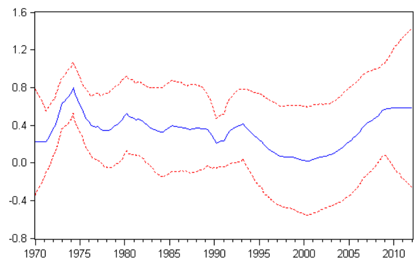
(b) Belgium



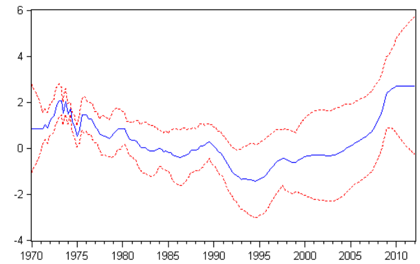
(c) Canada



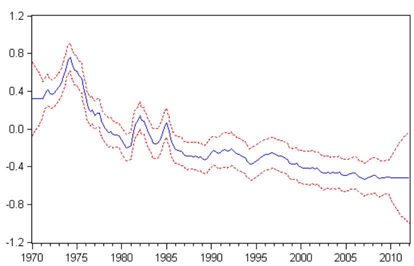
(d) Finland



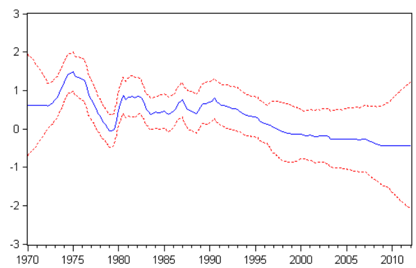
(e) Germany



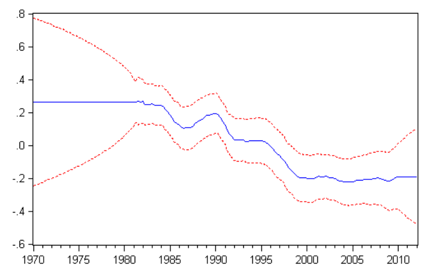
(f) Japan



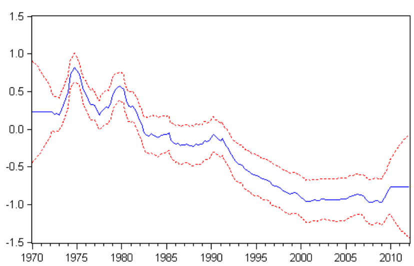
(g) Netherlands



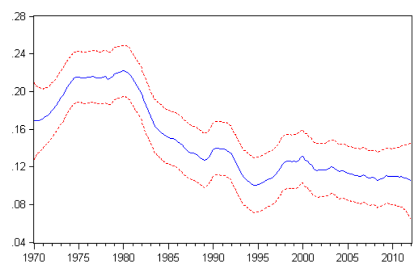
(h) Norway



(i) Sweden

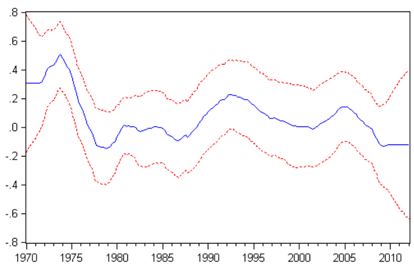


(j) UK

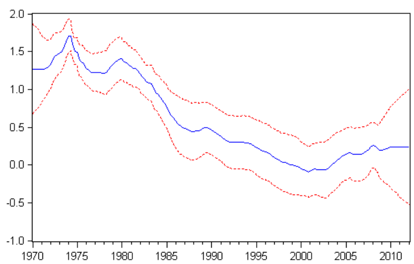


(k) US

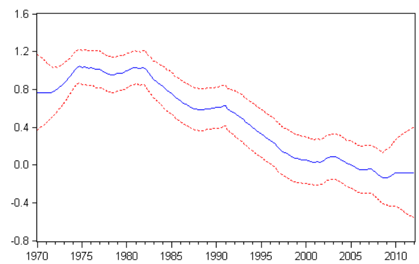
Figure 8: Smoothed estimates for γ_t: Model (21b)



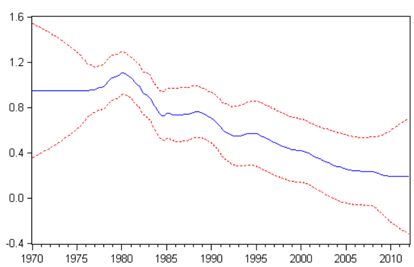
(a) Austria



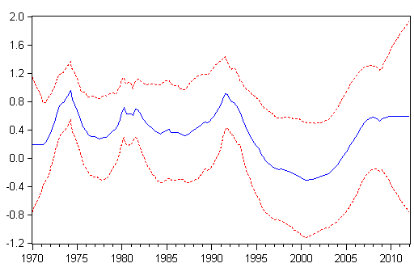
(b) Belgium



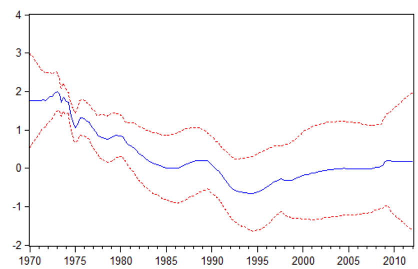
(c) Canada



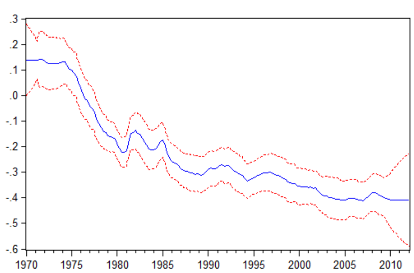
(d) Finland



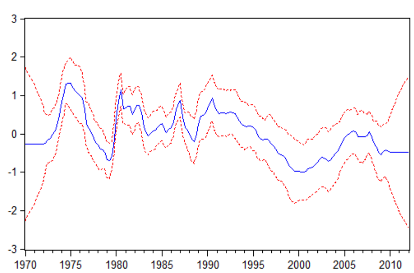
(e) Germany



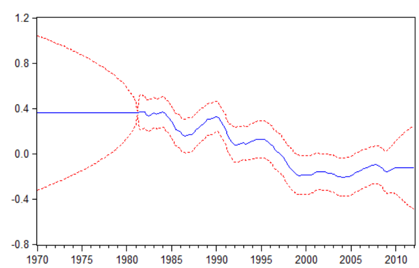
(f) Japan



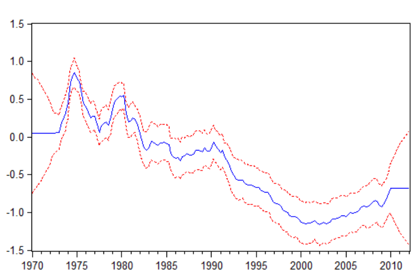
(g) Netherlands



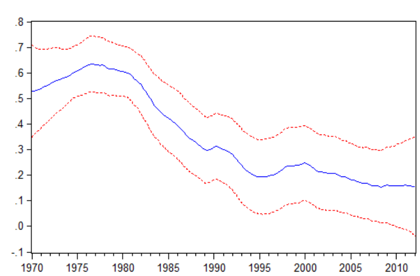
(h) Norway



(i) Sweden



(j) UK



(k) US

Figure 9: Correlations between unemployment (y $axis$) and wage inflation (x $axis$)