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Abstract. This paper uses data for the UK and the Netherlands (1983q4-2011q4) to test if hysteresis occurs in these economies, and through what mechanisms. The novelty of the paper resides in the use of a VAR-IRF that encompasses previous hysteresis studies and the use of specific Labour Market Institutions shocks. This allows us to disentangle what specific demand and supply-variables affect unemployment in the long-run. Further we also investigate the impact of different supply and demand-shock on long-term unemployment. Our findings suggest that there is hysteresis in both countries, and that it happens through different channels, namely, long-term unemployment, productivity, capital stock and real long-term interest rates. These results have implications for structural and macroeconomic policies that we also discuss.

JEL Classification
C32, E24, J64

Keywords
Hysteresis, NAIRU, Labour Market Institutions, Impulse-Response analysis.
1. Introduction

It is usually argued that changes in Labour Market Institutions (LMI) can explain the evolution of unemployment in the UK and the Netherlands since the 1980s (Nickell 1998, Nickell and Van Ours 2000, Broersma et al. 2000). However, the persistence of long-term unemployment and the impact of recessions on investment over the same period, see Fig. 1, suggest that hysteresis also influenced unemployment in these economies. This paper examines this hypothesis, we test if there is hysteresis in the UK and the Netherlands and what demand variables cause such phenomenon.

Several approaches have been used to test the hysteresis hypothesis in country-specific studies. Some authors apply unit root and stationarity tests to study the properties of unemployment series. If unemployment is found to be mean reverting it is interpreted as evidence in favour of the NAIRU a la Layard et al. (1991, Ch.8). Alternatively, if unemployment exhibits a unit root it is taken as evidence of hysteresis. Some recent reviews of this literature can be found in Romero-Avila and Usabiaga (2008) or Fosten and Ghoshray (2011). Evidence for the UK and the Netherlands is mixed and sensitive to the inclusion of structural breaks and changes in the sample period. Furthermore, this approach cannot disentangle what demand factors, e.g. long-term unemployment or capital stock, propitiate unit roots.

This problem does not arise in other branches of the literature that aim at testing specific hysteresis mechanisms rather than the properties of unemployment series. This is case of the wage-equation literature that examines the response of real wages to unemployment duration. This serves to test Blanchard and Summers (1986) hypothesis, that shocks that increase the share of long-term unemployment increase wage pressure and in turn the NAIRU, creating hysteresis. Available evidence for the UK is mixed, Nickell (1987), Manning (1994) and Arulampalam et al. (2000) find support for this hypothesis, but this is disputed by Blanchflower and Oswald (1994) and Bell and Blanchflower (2014). In the Netherlands, this hypothesis is rejected by (Graafland 1991, 1992). However, this literature focuses on the wage-equation and does not control for demand
variables that affect the price-setting behaviour of firms, which could also cause hysteresis, such as, capital stock (Bean 1989, Arestis and Sawyer 2005) or interest rates (Fitoussi and Phelps 1988, p.57, Blanchard 2002). These omissions are problematic because they can bias results.

Other authors use time-varying estimation methods to identify the structural breaks on unemployment and the variables that cause such shifts, e.g. Logeay and Tober (2006) and Srinivasan and Mitra (2014) for Germany and the UK, respectively. Srinivasan and Mitra find no evidence of hysteresis in the UK, although they only use labour-supply measures to control for hysteresis. Further, this single-equation approach does not take account of the interactions that shocks generate in the labour market.

To control for these interactions researchers use VAR-systems to model the labour market and then use the associated Impulse Response Functions (IRF) to evaluate the impact of shocks. Carstensen and Hansen (2000), Gambetti and Pistoresi (2004) and Binotti and Ghiani (2008) apply this approach to Germany and Italy\(^1\). They simulate labour-supply shocks to tests Blanchard and Summers hypothesis, and productivity shocks to evaluate claims that this variable can reduce unemployment permanently (Stiglitz 1997, Ball and Mankiw 2002). However, these studies might also be subject to biases as they overlook the potential influence of capital stock and real long-term interest rates on unemployment.

Our paper, extends available literature by estimating a VAR-IRF that allows us to test for hysteresis effects through unemployment duration, productivity, capital stock and real long-term interest rates. This allows us to disentangle what specific demand-factors affect unemployment in the long-run while controlling for the interactions in the labour market and avoiding potential biases that existing literature might be subject to. We use the most recent data for the UK and the Netherlands to estimate a Cointegrated-VAR model for each country, and simulate different demand-shocks, using the associated IRFs, to test different hysteresis hypothesis. Further, we also consider the impact of specific LMI-shocks to investigate the LMI-unemployment link. This is in contrast to existing VAR-IRF literature, where this link is studied by simulating generic wage-shocks. Our approach seems more adequate for policy design given that available evidence suggests that not all LMI increase the NAIRU (Nickell 1997, OECD 2006, Layard and Nickell 2011, Ch.7). Finally, exploiting the system nature of our estimation we also study the impact of different supply and demand-shocks on long-term unemployment.

The paper proceeds as follows. Section 2 presents our theoretical model. Section 3 explains our methodology. Section 4 examines the data. Section 5 presents our estimations. Section 6 presents our IRF-simulations. Section 6 summarizes our findings and their implications.

\(^1\) See also Dolado and Jimeno (1997) or Fabiani et al. (2001).
2. A wage and price-setting model

Our theoretical model draws from the well-known NAIRU model presented in Layard et al. (1991, Ch. 8) and Layard and Nickell (2011, Ch.1-2), LNJ hereafter. Although we extend this model to account for the following hysteresis factors, long-term unemployment, productivity, capital stock and real long-term interest rates. We assume that the labour and goods market operate under imperfect competition, which gives workers and firms price-making power. In the labour market, imperfect competition is usually characterized by the presence of unions or efficiency wages. Accordingly, workers try to achieve a “target” real wage \((w - p)\) here denoted by the following equation when expectations are fulfilled (all the variables are in logarithm):

\[
w - p = -\omega_1 u + \omega_2 (y - l) + \omega_4 grr + \omega_5 t^w + \omega_6 up + \omega_7 lu \tag{1}
\]

Equation (1) is a negative function of unemployment \((u)\), which captures the state of the labour market, and a positive function of productivity \((y - l)\) that encourages workers income claims. As in LNJ’s model, workers’ “Target” is also pushed-up by several LMIs, viz-a-viz, unemployment benefits \((grr)\), the tax-wedge \((t^w)\) and unions’ power \((up)\). We do not consider other LMIs, such as, employment legislation or minimum wages because existing evidence is generally not supportive of their influence on unemployment (OECD 2006, p.59-107, Layard and Nickell 2011, Ch.7). Further, equation (1) also grows with long-term unemployment \((lu)\) reflecting Blanchard and Summers (1986) hypothesis, BS hereafter, that increases in unemployment duration rises insiders bargaining power and wages, which in turn rises the NAIRU. See also Ball (1999, 2009) or Krueger et al. (2014).

In the goods market, imperfect competition takes the form of monopolistic or oligopolistic competition. This allows firms to set prices as a mark-up over wages \((p - w)\) moderated by unemployment and productivity, hence, the term “feasible” real wage to refer to this mark-up, here illustrated by equation (2) (with fulfilled expectations).

\[
p - w = -\varphi_1 u - \varphi_2 (y - l) - \varphi_3 (i - \Delta p) \tag{2}
\]

Equation (2) is also a negative function of capital stock \((k)\) to account for hysteresis through capital-scrapping. Bean (1989) and Arestis and Sawyer (2005) argue that after a negative shock, capital accumulation slows-down scrapping part of the productive capacity of the economy. As a result, some of the workers that lost their jobs cannot be re-employed and add-up to the existing unemployment equilibrium, i.e. the NAIRU. Further, equation (2) is a positive function of real long-term interest rates \((i - \Delta p)\) to reflect claims by Fitoussi and Phelps (1988, p.57) and Blanchard (2002), that firms’ (real) long-term cost of borrowing increases firms mark-up, and in turn the NAIRU.

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2 Similar models can be found in Manning (1993), Nickell (1998) or Gianella et al. (2008).
3 See the surveys in Layard et al. (1991, Ch.2-4) or Manning (2011).
4 See Sawyer (1982), Blanchard (1988) or Layard et al. (1991,Ch. 7).
5 See also Rowthorn (1995, 1999).
Equating (1) and (2) to solve for unemployment, we find equation (3), the unemployment level that makes workers and firms income claims compatible, i.e. the NAIRU. Further, solving for $w - p$, we find the long-run real wage equilibrium associated with this NAIRU, equation (4).

$$u^* = \beta_{12}(y - l) + \beta_{14}lu + \beta_{15}grr + \beta_{16}t^w + \beta_{17}up + \beta_{18}k + \beta_{19}(i - \Delta p) \quad (3)$$

$$\beta_{12} = \omega_2 - \varphi_3, \quad \beta_{14} = \frac{\omega_2}{\omega_1 + \varphi_1}, \quad \beta_{15} = \frac{\omega_4}{\omega_1 + \varphi_1}, \quad \beta_{16} = \frac{\omega_5}{\omega_1 + \varphi_1}, \quad \beta_{17} = \frac{\omega_6}{\omega_1 + \varphi_1}, \quad \beta_{18} = -\frac{\varphi_2}{\omega_1 + \varphi_1}, \quad \beta_{19} = \frac{\varphi_5}{\omega_1 + \varphi_1}$$

$$(w - p)^* = \beta_{22}(y - l) + \beta_{24}lu + \beta_{25}grr + \beta_{26}t^w + \beta_{27}up + \beta_{28}k + \beta_{29}(i - \Delta p) \quad (4)$$

$$\beta_{22} = \left(\omega_2 - \omega_1 \frac{\omega_2 - \varphi_3}{\omega_1 + \varphi_1}\right), \quad \beta_{24} = \varphi_1 \frac{\omega_1}{\omega_1 + \varphi_1}, \quad \beta_{25} = \varphi_1 \frac{\omega_4}{\omega_1 + \varphi_1}, \quad \beta_{26} = \varphi_1 \frac{\omega_5}{\omega_1 + \varphi_1}, \quad \beta_{27} = \varphi_1 \frac{\omega_6}{\omega_1 + \varphi_1}, \quad \beta_{28} = \omega_1 \frac{\varphi_2}{\omega_1 + \varphi_1}, \quad \beta_{29} = -\omega_1 \frac{\varphi_5}{\omega_1 + \varphi_1}$$

The NAIRU described by equation (3) is a function of LMI and demand variables. If $\beta_{12} = \beta_{14} = \beta_{18} = \beta_{19} = 0$, the NAIRU is exclusively determined by LMI, as argued by LNJ. In our case, $grr, t^w$ and $up$. These $\beta$ restrictions imply the following restrictions in equations (1) and (2). First, $\beta_{12} = 0$ meaning that the NAIRU is neutral to productivity, requires $\omega_2 = \varphi_3$, i.e. productivity gains are fully reflected in workers real wages. However, if wages are slow to react to changes in productivity, i.e. $\omega_2 < \varphi_3$ for long lapses of time, productivity reduces the NAIRU ($\beta_{12} < 0$) as argued by Stiglitz (1997) and Ball and Mankiw (2002).  

Second, $\beta_{14} = 0$ requires $\omega_{11} = 0$, meaning that there is no hysteresis through unemployment duration. However, if $\omega_{11} > 0$, then greater long-term unemployment increases wage claims and the NAIRU ($\beta_{14} > 0$) as proposed by BS. Third, $\beta_{18} = 0$ requires $\varphi_2 = 0$, ruling out capital-scraping. However, if $\varphi_2 > 0$, then more productive capacity reduces firms mark-up and the NAIRU ($\beta_{18} < 0$) as advocated by the capital-scraping hypothesis. Fourth, $\beta_{19} = 0$ requires $\varphi_5 = 0$, the cost of borrowing does not affect the “feasible” real wage, either the NAIRU. However, if $\varphi_5 > 0$ then real long-term interest rates rises firms’ mark-up and in turn the NAIRU ($\beta_{19} > 0$).

Hence, if any of the restrictions $\beta_{12} = \beta_{14} = \beta_{18} = \beta_{19} = 0$ does not hold, the NAIRU is determined by either $y - l$ or $lu$ or $k$ or $i - \Delta p$. These variables are sensitive to the evolution of demand (and macroeconomic policy), thus, there will be hysteresis.

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6 Other authors suggest that productivity can also affect the NAIRU depending on its relative value to wealth (Phelps 2000) and interest rates (Pissarides 2000).

7 For some authors the interest rates-NAIRU link does not create a monetary policy-NAIRU link (Hian Teck and Phelps 1992, p. 896, Gianella et al. 2008, p. 21). They argue that changes in real long-term interest rates are not the result of changes in monetary policy but of the evolution of financial markets along with governments' fiscal position. This claim is at odds with Central Bank pass-through estimates in the UK (Gulmaraes 2012) and the Netherlands (ECB 2009). Hence, we treat real long-term interest rates as a demand variable.
3. Methodology

A VAR-IRF approach is particularly well equipped to study hysteresis in the labour market (Dolado and Jimeno 1997, Fabiani et al. 2001). IRF-simulations allow us to simulate different demand-shocks to investigate what variables have a long-term impact on unemployment, i.e. affect the NAIRU and cause hysteresis, and what variables only have temporary or cyclical influence. Further, the system nature of the VAR-IRF provides a wealth of estimates capturing not only the influence of a specific variable on unemployment, e.g. productivity, but also on the rest of variables of the system, e.g. real wages, to assess the source of the relationship with unemployment. That is, it allows us to test the restrictions on \( \beta \), together with those for \( \omega \) and \( \varphi \). Finally, the VAR accommodates the potential endogeneity between our variables.

We estimate a VAR for the vector \( z_t \) that contains all the variables included in equations (1)-(4), i.e. \( z_t = (w_t - p_t, y_t - l_t, u_t, lu_t, grr_t, t^w, up_t, k_t, i_t - \Delta p_t)' \). Unit root and stationarity tests suggest that we should treat these variables as \( I(1) \), see Appendix 3. Hence, we adopt a cointegrated-VAR approach illustrated by the following Vector Error Correction Model (VECM):

\[
\Delta z_t = c_0 + \Phi_1 \Delta z_{t-1} + \cdots + \Phi_{n-1} \Delta z_{t-n+1} + T \gamma \beta' \left[ \frac{T-1}{T} \right] + \lambda x_t + \epsilon_t \quad (6)
\]

Where \( \Delta \) is the first difference operator, \( z_t \) is a \((9 \times 1)\) vector for our \( I(1) \) variables. \( c_0 \) is a vector of intercepts. \( \Delta z_{t-n+1} \) is the higher lag of \( \Delta z_t \) considered, where \( n \) denotes the lag order of the underlying VAR. \( \gamma \beta' \left[ \frac{T-1}{T} \right] \) is the Error Correction Mechanism (ECM) of the system, with \( \beta' \) and \( \gamma \) denoting the matrix of long-run and loading coefficients, respectively. \( T \) is a vector of time-trends that account for the deterministic trends that some variables exhibit, see Figs. 2-6, although \( T \) is restricted to the ECM-term to avoid quadratic trends. \( x_t \) is a \((h \times 1)\) vector of exogenous \( I(0) \) variables that controls for past external cost-shocks with two lags of \( \Delta p^m_t \). In the Dutch case, \( x_t \) also contains the dummy \( D95q2 \) that accommodates an outlier, which caused serial correlation in preliminary estimations. \( \epsilon_t \) is a vector of error terms, Normally and Independently Distributed (NID).

Drawing from AIC and SBC selection criteria we favour a lag order \( n = 2 \). This choice and the composition of \( x_t \) are the result of experimenting with several specifications, until a parsimonious but informative lag structure that provides satisfactory diagnostic test results is found. Hence, our empirical specification is the following:

\[
\begin{pmatrix}
\Delta(w - p)_t \\
\Delta(y - l)_t \\
\Delta u_t \\
\Delta lu_t \\
\Delta grr_t \\
\Delta t^w \\
\Delta up_t \\
\Delta k_t \\
\Delta(i - \Delta p)_t \\
\end{pmatrix}
= \begin{pmatrix}
c_{0, w-p} \\
c_{0, y-l} \\
c_{0, u} \\
c_{0, lu} \\
c_{0, grr} \\
c_{0, tw} \\
c_{0, up} \\
c_{0, k} \\
c_{0, i-\Delta p} \\
\end{pmatrix}
+ \Phi_1
\begin{pmatrix}
\Delta(w - p)_{t-1} \\
\Delta(y - l)_{t-1} \\
\Delta u_{t-1} \\
\Delta lu_{t-1} \\
\Delta grr_{t-1} \\
\Delta t^w_{t-1} \\
\Delta up_{t-1} \\
\Delta k_{t-1} \\
\Delta(i - \Delta p)_{t-1} \\
\end{pmatrix}
+ \gamma \beta' \left[ \frac{T-1}{T} \right] + \lambda x_t + 
\begin{pmatrix}
\epsilon_{w-p} \\
\epsilon_{y-l} \\
\epsilon_u \\
\epsilon_{lu} \\
\epsilon_{grr} \\
\epsilon_{tw} \\
\epsilon_{up} \\
\epsilon_k \\
\epsilon_{i-\Delta p} \\
\end{pmatrix} \quad (7)
\]

Where, \( x_{t, UK} = (\Delta p^m_{t-1}, \Delta p^m_{t-2})' \); \( x_{t, Neth} = (\Delta p^m_{t-1}, \Delta p^m_{t-2}, D95q2)' \).
Our empirical strategy proceeds as follows. First, we test for cointegration in $z_t$ using the Maximum eigenvalue ($\lambda_{\text{max}}$) and Trace ($\lambda_{\text{trace}}$) tests. Second, we identify the long-run relationships that exist among our variables using restrictions drawn from economic theory, as suggested by Pesaran and Shin (2002) and Garrat et al. (2006, Chapter 6). This provides a first test of the $\beta$-restrictions by identifying what variables have a long-run relationship with unemployment. Third, we estimate the short-run dynamics of the system, evaluate the goodness of the fit and the stability of the estimated VAR. Fourth, we use IRF to simulate different demand and supply-shocks (of one standard deviation of the residuals, $\sigma_e$) and test the hysteresis hypothesis embedded in our model. It should be noted that all IRF reported in section 6 are Generalized-IRF (GIRF), to ensure that the ordering of variables in the VAR-system does not affect the outcome of our simulations (Garrat et al. 2006, p.142).

These are the IRF-simulations presented below and their correspondence with our theoretical model: 

i) Shocks in unemployment ($\sigma_{\Delta u}$) and long-term unemployment ($\sigma_{\Delta tl_u}$) to assess the response of real wages to an increase in unemployment duration. This tests $\omega_{11} = 0$ and $\beta_{14} = 0$.

ii) Productivity shock ($\sigma_{\Delta y-l}$) to evaluate its impact on unemployment and real wages, i.e. $\beta_{12} = 0$, $\omega_2 = \varphi_3$.

iii) Capital stock shock ($\sigma_{\Delta k}$) to assess its impact on unemployment, this tests $\beta_{18} = 0$.

iv) Shock in real long-term interest rates ($\sigma_{\Delta(\Delta r)}$) to evaluate the response of unemployment, i.e. $\beta_{19} = 0$.

v) Three LMI-shocks, unemployment benefits ($\sigma_{\Delta gr}$), labour taxation ($\sigma_{\Delta tw}$), and unions power ($\sigma_{\Delta up}$) to assess the response of unemployment and draw policy implications for labour market reforms. These tests, $\beta_{15} > 0$, $\beta_{16} > 0$ and $\beta_{17} > 0$.

vi) We examine the response of long-term unemployment to the above mentioned demand and supply-shocks to draw further policy conclusions.

4. Data

Our dataset contains quarterly data for the UK and the Netherlands from 1983q4 to 2011q4. The sample period is determined by data availability. The main source of data is OECD’s statistical office, although we also employ data from UK’s Office of National Statistics (ONS) and the IMF, see Appendix 1 for further details. Before presenting our estimations we examine our data in the following figures (in logarithm scale). Fig. 2 shows the evolution of unemployment. In both countries, unemployment seems to trend downwards after peaking in the early-1980s, and despite hikes in the early-1990s, early-2000s, and after 2007.
Fig. 2. Unemployment

Fig. 3 shows the evolution of our LMIs (solid line) against unemployment (dotted line). Unemployment benefits ($grr$), panels (a) and (b), the tax-wedge ($t^w$), panels (c) and (d), and unions’ power ($up$), panels (e) and (f), all trend downwards for most of the sample period. This reflects the labour market reforms introduced in both countries since the 1980s (Siebert 1997, Nickell and Van Ours 2000, Brandt et al. 2005, OECD 2000, 2006). The downward trends of LMI and unemployment since the 1980s might suggest that there is a positive relationship between these variables. However, close examination shows that only unemployment benefits, particularly in the UK, moved with unemployment over the sample period. The tax-wedge, panels (c) and (d), moves downwards with unemployment up to the late-1980s, but for most of the 1990s and 2000s $t^w$ and unemployment move in opposite directions. Similarly, unions’ power shown in panels (e) and (f). Our IRFs will confirm that only unemployment benefits have a positive long-run relationship with unemployment.
Fig. 3. Labour Market Institutions

We turn now to demand-factors that could cause hysteresis. Fig. 4, panels (a) and (b), present the evolution of long-term unemployment \((lu)\) against total unemployment \((u)\). In both countries \(lu\) mirrors the behaviour of overall unemployment with some delay. This suggests that demand shocks can increase (and reduce) long-term unemployment in line with BS hypothesis. In panels (c) and (d), we observe the evolution of capital accumulation \((\Delta k)\) along with unemployment \((u)\). In both countries, there seems to be a negative relationship between \(\Delta k\) and \(u\), as periods of greater accumulation coincide with reductions of unemployment, e.g. the second half of the 1980s and 1990s. And periods of lower investment come with rising unemployment, e.g. the early-1980s, early-1990s and after 2007. This behaviour is consistent with the capital-scrapping hypothesis. Our IRFs suggest that this relationship is statistically significant.
Fig. 4. Long-term unemployment and capital accumulation

Fig. 5, panels (a) and (b), compare the evolution of productivity ($y - l$) and unemployment ($u$). The former exhibits a clear upward trend in both countries, despite marked falls in 2007, and several periods of stagnation in both countries in the early-1980s, between 1988-1992, early-2000s, and in the UK, after 2008. Some of these slowdowns in $y - l$ coincide with hikes in unemployment, suggesting that there is a negative relationship between these variables, which our IRFs will confirm. Looking at panel (c), this relationship could be the result of real wages lagging behind productivity in the UK, as the reductions in unemployment during the late 1980s, and from mid-1990s to 2007, are characterized by $y - l$ growing faster than $w - p$. This is unclear in the Dutch case, panel (d), where $w - p$ grow above $y - l$ in periods of rising but also falling unemployment.
Finally, Fig. 6 compares the evolution of real long-term interest rates \((i - \Delta p)\) and unemployment. In both countries, there is an initial period of relatively high but stable interest rates until the early-1990s, period in which the highest levels of unemployment were recorded. The rest of the sample period is characterized by an intense fall in \(i - \Delta p\) that coincides with falling unemployment from the mid-1990s to the early-2000s, but also with rising joblessness in the 2000s. Hence, from these figures it is unclear whether there is a relationship between these variables and whether this is positive or negative.

**5. Model estimation**

We start our empirical analysis by testing for cointegration in 
\[ z_t = (w_t - p_t, y_t - l_t, u_t, l_u_t, g_{rr_t}, t_t, i_t, u_t, i_t - \Delta p_t)' . \]
Table 1 presents our results. In both countries, Maximum eigenvalue \(\lambda_{max}\) and Trace \(\lambda_{trace}\) tests support the existence of cointegration, as the null \(r = 0\) is rejected at one percent significance level for both tests\(^8\). However, tests differ with regard to the number of long-run relationships. In the UK’s case, \(\lambda_{max}\) fails to reject the null hypothesis of having two long-run relationships, while \(\lambda_{trace}\) fails to reject having four cointegrated vectors, both at one percent. For the Netherlands, \(\lambda_{max}\) and \(\lambda_{trace}\), fail to reject the null of having two and five cointegrated vectors, respectively. Weighting these results against the predictions from our theoretical model, which suggest that there are two cointegrated vectors among our variables, equations (3) and (4), it seems reasonable to proceed under the assumption of \(r = 2\). This choice is vindicated by the diagnostic tests for our short-run equations, see Table 3 below, which suggest that the model is specified satisfactorily.

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\(^8\) Given the well-known size problems of these tests we use 1\% critical values rather than standard 5\%.
Cointegration tests

Table 1. Cointegration tests

|       | UK          |             | Netherlands |             |
|-------|-------------|-------------|-------------|-------------|
|       | $H_0$      | $H_1$      | $\lambda_{\text{max}}$ test | $\lambda_{\text{trace}}$ test | $\lambda_{\text{max}}$ test | $\lambda_{\text{trace}}$ test |
| $r=0$ | $r=1$      | 68.28 [0.009] | 318.62 [0.000] | 80.18 [0.000] | 336.11 [0.000] |             |
| $r=1$ | $r=2$      | 63.12 [0.006] | 250.34 [0.000] | 62.02 [0.009] | 255.92 [0.000] |             |
| $r=2$ | $r=3$      | 54.82 [0.013] | 187.22 [0.000] | 49.76 [0.057] | 193.90 [0.000] |             |
| $r=3$ | $r=4$      | 37.05 [0.266] | 132.40 [0.003] | 40.88 [0.117] | 144.14 [0.000] |             |
| $r=4$ | $r=5$      | 35.64 [0.007] | 95.35 [0.013] | 36.60 [0.075] | 103.25 [0.002] |             |
| $r=5$ | $r=6$      | 22.34 [0.481] | 59.72 [0.105] | 30.45 [0.076] | 66.65 [0.026] |             |
| $r=6$ | $r=7$      | 18.91 [0.235] | 37.37 [0.161] | 16.29 [0.053] | 36.20 [0.201] |             |
| $r=7$ | $r=8$      | 14.36 [0.239] | 18.46 [0.320] | 14.71 [0.217] | 19.92 [0.234] |             |
| $r=8$ | $r=9$      | 4.10 [0.729]  | 4.10 [0.727]  | 5.21 [0.575]  | 5.21 [0.574]  |             |

Note: $p$-values in square brackets. Test statistics obtained from applying $\lambda_{\text{max}}$ and $\lambda_{\text{trace}}$ tests to $z_t$, using a VAR(2) with unrestricted intercepts, restricted trend-coefficients and two lags of $\Delta p^m$. Netherlands’ computations also includes $0.95q^2$. For both countries we use 110 observations, 1984q3-2011q4.

Next, we study what variables take part in these two cointegrated vectors by identifying the matrix of long-run coefficients, $\beta$ in equation (7). For this purpose, we draw from economic theory and experiment with several schemes. These include $\beta$’s exclusively determined by LMIs and combinations of these with different versions of $\beta$ allowing for demand variables, see Appendix 2 for further details. In both countries, there is little support for a $\beta$ exclusively determined by LMIs, as restrictions excluding all demand-factors are insignificant for both countries. In the UK, the Likelihood Ratio ($LR$) statistic is $X^2_{LR(10)}=52.351$, and in the Netherlands $X^2_{LR(10)}=56.647$. In both cases, $p$-value=0.000.

Further, excluding $l_u$ and $k_t$ from the unemployment vector pushes $\beta$ into rejection in both countries, suggesting that these variables are cointegrated with unemployment. Unit-proportionality between real wages and productivity seems supported in the UK but not in the Netherlands. Further, the co-trending hypothesis (excluding $T$ from the cointegrated vector) seems to hold in both vectors for the UK, but only in the unemployment vector for the Netherlands. Using this information we experiment until we find a significant $\beta$-matrix of long-run coefficients for each country that we adopt as our preferred long-run specifications, here reported in Table 2:

Table 2. $\beta$ matrix of long-run elasticities

| United Kingdom |            |            | Netherlands |            |
|----------------|------------|------------|-------------|------------|
| $(w-p)_t$      | $(y-l)_t$  | $u_t$      | $l_u$       | $g_{r_t}$  | $\tau^p_t$  | $u_p$       | $k_t$       | $(1-\Delta p)_t$ | $T$          |
| 0.000          | 0.000      | -1.000     | 0.853      | 0.000      | 0.000       | 0.000       | -0.147     | (0.073)       | 0.000         | 0.000         |
| -1.000         | 1.000      | 0.000      | 0.254      | 0.000      | -1.375      | 0.026       | 0.000      | -0.136       | (0.042)       | 0.000         |
| Netherlands    |            |            |            |            |            |            |            |            |            |
| $(w-p)_t$      | $(y-l)_t$  | $u_t$      | $l_u$       | $g_{r_t}$  | $\tau^p_t$  | $u_p$       | $k_t$       | $(1-\Delta p)_t$ | $T$          |
| 0.000          | 0.000      | -1.000     | 2.565      | 0.000      | -5.227      | 0.000       | -2.928     | (0.039)       | -1.594       | 0.000         |
| -1.000         | 0.300      | 0.000      | 0.000      | -0.101     | 0.000       | -0.273      | 0.484      | -0.054       | (0.005)       | -0.005       |

Note: Asymptotic Standard Errors for $\beta$-coefficients in brackets. $^3$ indicates significant at 5% and $^*$ at 10%. $X^2_{LR}(q-r^2)$ is chi-square statistic for $LR$-test. $p$-values for $LR$-test in square brackets. $q$-number over-identifying restrictions, $r^2$ = just – identified restrictions. $l_u$ and $l_{u*}$ are maximum value of log-likelihood function under $q$ and $r^2$, respectively. We adopt $r^2 = 4$: First vector, $u_t$ normalized and $(w-p)$, excluded to identify (3), whereas, second vector, $(w-p)$, normalized and $u_t$ excluded to identify (4). For both countries we use 110 observations, 1984q3-2011q4.
According to Table 2, in the UK unemployment is cointegrated with long-term unemployment and capital stock, with long-run elasticities of 0.83 and -0.15, respectively. These findings are in accordance with previous cointegration analysis of British unemployment (Arestis and Biefang-Frisancho Mariscal 1998, 2000). Results for the Netherlands are similar, an increase in \( lu \) of 1% increases the overall unemployment rate by 2.05% in the long-run, whereas, an increase in \( k \) of one percent reduces unemployment by 2.93%. Arestis et al. (2007) report similar results. Further, in the Netherlands, we find unemployment cointegrated with real long-term interest rates, with an elasticity of -1.59. This sign is unexpected because the cost of borrowing and unemployment are usually thought to have a positive relationship. We investigate this issue below.

Finally, in the second cointegrated vector, real wages are positively cointegrated with productivity in both countries, although the intensity of this relationship varies, as suggested by Fig. 5 (c)-(d). In the UK, we find a long-run one-to-one relationship between real wages and productivity as also reported by Arestis and Biefang-Frisancho Mariscal (1998). While, in the Netherlands, the long-run elasticity of real wages to productivity is only 0.3, in contrast to Schreiber (2012) who find long-run unit-proportionality between these variables. It should be noted that these cointegrated vectors are clearly significant. In the UK, the LR statistic is \( X_{LR}^2(10) = 10.307 \) with \( p \)-value = 0.414, and in the Netherlands, \( X_{LR}^2(7) = 9.532 \) with \( p \)-value = 0.217. Further, they provide evidence of long-run links between unemployment and demand-factors that our IRF will confirm.

Next, we estimate the short-run dynamics of the system and evaluate our VAR-model. Table 3 presents our estimates for the ECM-terms and diagnostic tests for unemployment short-run equations, \( \Delta u_t \) in (7), see appendix 2 for further details. The \( \Delta u_t \) regressions for both countries pass all the diagnostic tests at the standard 5% significance level\(^9\), this suggests that our model is specified satisfactorily. The adjusted \( R \)-squares are also reasonably high, 0.52 for the UK, and 0.77 for the Netherlands.

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\(^9\) Netherlands’ regression fails to pass the normality test, although this only invalidates inference using the \( t \)-test.
Table 3. Short-run dynamics

|        | UK 1984q3-2011q4 | Netherlands 1984q3-2011q4 |
|--------|------------------|---------------------------|
|        | Δu_t             | Δu_t                      |
| $\xi_{1t-1}$ | -0.022 [0.438] | -0.031$^*_{[0.009]}$     |
| $\xi_{2t-1}$ | 0.211$^*_{[0.004]}$ | 0.482 [0.156]             |

Observations: 110 110
Adj. R$^2$: 0.522 0.768

$X^2_{SC}(4)$: 8.452$^*$ [0.076] 3.817 [0.431]
$X^2_{FF}(1)$: 2.754$^*$ [0.097] 0.048 [0.825]
$X^2_{Norm}(2)$: 4.354 [0.113] 12.45$^*$ [0.002]
$X^2_{Het}(1)$: 0.001 [0.975] 0.108 [0.743]

Note: p-values for t-tests and diagnostic tests in square brackets. $^*$ indicates significant at 5% and $^*$ at 10%. Adj. R$^2$=Adjusted R-square. $X^2_{SC}, X^2_{FF}, X^2_{Norm}, X^2_{Het}$ are chi-square statistics for Serial correlation (SC), Functional Form (FF), Residuals Normality (NORM) and Heteroscedasticity (HET) tests, respectively.

For the UK: $\xi_{1t} = u - 0.832u + 0.147k$
$\xi_{2t} = (w - p) - (y - f) - 0.253u + 1.375t_0 + 0.026u + 0.135(t - Δp)$
Netherlands: $\xi_{1t} = u - 2.054u + 5.226t_0 + 2.928k + 1.594(t - Δp)$
$\xi_{2t} = (w - p) - 0.300(y - f) + 0.101gr + 0.273up - 0.484k$
+0.054(t - Δp) + 0.0057

The goodness of the fit is further illustrated by Fig. 7 that compares changes in unemployment implied by our estimations against historical data. Overall, our fitted values seem to do a good job in describing unemployment fluctuations in both countries. Further, the eigenvalues of the companion matrix, see Appendix 2, are all within the unit circle, which suggests that our VAR-system is stable and adequately specified. Thus, we conclude that our estimations capture satisfactorily the long-run and short-run properties of unemployment in both countries and proceed with our IRF-simulations.

![Fig. 7. Actual and fitted Δu_t](image)
6. Impulse-response analysis

6.1 Hysteresis hypothesis

We start by testing BS hypothesis that long-term unemployment increases real wages and in turn the NAIRU ($\omega_{11} > 0, \beta_{14} > 0$). For this purpose, Fig. 8 shows the response of real wages ($w - p$) to a shock on overall unemployment ($u$) and to a rise on long-term unemployment ($lu$). Panels (a) and (b), show that a rise in $u$, moderates real wages by about 5% and 1.5%, in the UK and the Netherlands, respectively. On the other hand, a rise in $lu$, panels (c) and (d), has no significant impact in the UK, but it increases $w - p$ in the Netherlands. This suggests that demand shocks that modify the share of long-term unemployment generate hysteresis in the Netherlands, i.e. $\omega_{11} > 0$ and $\beta_{14} > 0$, but not in the UK. Our results for the UK are in line with Blanchflower and Oswald (1994) and Bell and Blanchflower (2014), but our findings for the Netherlands are in contrast to Graafland (1991, 1992).

![Fig. 8. Unemployment duration-shock](image)

Note: Solid lines denote IRF point estimates and dashed lines 95% confidence interval (CI).

Fig. 9 evaluates the effect of a productivity shock on unemployment and real wages. Panels (a) and (b), show that, in both countries, a rise in productivity reduces unemployment permanently, i.e. reduces the NAIRU. In the UK, the unemployment rate falls on impact until it stabilizes, eight quarters after the shock, at a rate 17% lower than its benchmark. In the Netherlands, unemployment stabilizes at a rate approximately 20% lower than its baseline. Hence, Fig. 9, suggests that $\beta_{12} < 0$ for both countries. Previous results are mixed, but our findings reinforce evidence of a negative relationship between productivity and the NAIRU in both countries (Broersma et al. 2000, Hatton 2007).
According to Stiglitz (1997) and Ball and Mankiw (2002) the productivity-NAIRU link is due to the slowness of workers to adapt their income claims to changes in productivity, that is, \( \omega_2 < \varphi_3 \) for long lapses of time. Panels (c) and (d), test this by comparing the evolution of productivity \((y - l)\) and real wages \((w - p)\) after the above mentioned productivity-shock. In both countries, the initial response of \( w - p \) falls short to the rise in \( y - l \), further it is only after eight quarters in the UK, twelve in the Netherlands, that real wages start closing the gap. Hence, Fig. 8 suggests that productivity influences the NAIRU, and that this depends on how fast workers claims adapt to improvements in productivity conditions.

Figure 10, panels (a) and (b), present the impact of a capital stock shock on unemployment. In both countries, greater accumulation reduces unemployment permanently, providing support for \( \beta_{18} < 0 \) in both countries. In the UK, the unemployment rate falls for about 24 quarters until it stabilizes at a rate 17% lower than its baseline. Whereas, in the Netherlands, the shock pushes unemployment to a minimum after eight quarters and then stabilizes at a rate approximately 20% below its pre-shock level. These findings are in line with the capital-scraping hypothesis and reinforce available evidence of a negative long-run relationship between capital stock and unemployment in the UK and the Netherlands (Dreze and Bean 1990, Arestis and Biefang-Frisancho Mariscal 1998, 2000, Arestis et al. 2007).
According to the capital scrapping hypothesis, the link between capital stock and unemployment is bi-directional, i.e. increases in capital stock reduce unemployment, because firms scrap productive capacity during downturns. Panels (c) and (d), assess this claim by examining the response of capital stock to a rise in unemployment. We find that effectively, rises in unemployment slows-down capital accumulation permanently by 5% and 2.8% in the UK and the Netherlands, respectively.

Fig. 11, evaluates the link between unemployment and the cost of borrowing. Panel (a) shows that after a rise in real long-term interest rates, British unemployment falls to a rate approximately 11% below its benchmark. In the Netherlands, panel (b), the impact of this shock is more modest as unemployment stabilizes at a rate 2% below its initial value, and this fall is only marginally significant in the long-run. Hence, Fig. 11, suggests that $\beta_{19} < 0$ for both countries. This sign is unexpected because Fitoussi and Phelps (1988, p.57) and Blanchard (2002) predict a positive relationship and available panel evidence seems to support it, e.g. Gianella et al. (2008). However, these studies do not control for capital stock. Hence, a possible explanation for this sign discrepancy is that their positive real interest rate coefficient is in fact capturing the negative influence of capital stock over the NAIRU. This highlights the importance of controlling for different hysteresis-factors.
Honkapohja and Koskela (1999) and Bell-Kelton and Ballinger (2005) discuss different rationales for the negative influence of $i - \Delta p$ on unemployment. They argue that this sign could be the result of a wealth effect, by which higher real long-term interest rates rises funding available to firms rather than making it more expensive, i.e. $\varphi_5 < 0$ rather than positive. Alternatively, they argue that this negative sign could reflect the impact of the cost of borrowing on the opportunity cost of being unemployed\textsuperscript{10}. To investigate this possibility we examine the response of real wages to a shock in interest rates in panel (c) and (d). In the UK, $w - p$ increase permanently by 3% as a result of the rise in the cost of borrowing. This suggests that higher interest rates rise firms “Feasible” real wage, making lower unemployment at higher wages compatible, i.e. $\varphi_5 < 0$. In the Netherlands, the shock has no significant impact on workers income after 2-3 quarters, and although this can explain why the shock only has a marginal impact in the long-run, it does not explain the behaviour of unemployment in the short-run. Hence, further research might be necessary.

In sum, according to our Impulse-Response analysis, demand-shocks have a permanent effect on unemployment, i.e. there is hysteresis in the UK and the Netherlands.

6.2 Unemployment and LMI

Next, we investigate the LMI-unemployment link by simulating three supply-shocks, namely, unemployment benefits ($grr$), labour taxation ($t^w$) and unions’ power ($up$). Fig. 12 presents the response of unemployment to these shocks. An increase in $grr$, panels (a) and (b), rises unemployment permanently in both countries. In the UK, the unemployment rate increases on impact, until it stabilizes 20 quarters after the shock at a rate 20% above its baseline. While in the Netherlands, unemployment describes a J-curve, and after an initial fall it stabilizes at a rate 2% greater than its benchmark,

\textsuperscript{10} Our model can accommodate this by adding the following term to equation (1): $-\omega_7(i - \Delta p)$
although this rise is only marginally significant. Layard et al. (1991, p.441), Nickell and Bell (1995), Broersma et al. (2000) and Gianella et al. (2008) also find that unemployment benefits increase the NAIRU in both countries.

**UK**

(a) $u$ response to $grr$-shock ($\sigma_{grr} = 0.0031$)

(b) $u$ response to $grr$-shock ($\sigma_{grr} = 0.0070$)

(c) $u$ response to $t^w$-shock ($\sigma_{t^w} = 0.0036$)

(d) $u$ response to $t^w$-shock ($\sigma_{t^w} = 0.0113$)

(e) $u$ response to $up$-shock ($\sigma_{up} = 1.206$)

(f) $u$ response to $up$-shock ($\sigma_{up} = 0.0047$)

**Netherlands**

**Fig. 12. LMI shocks**

Note: Solid lines denote IRF point estimates and dashed lines 95% CI.

A shock in $t^w$, panels (c) and (d), results in permanent reductions of unemployment in both countries, although the timing of the adjustment varies. In the UK, unemployment stabilizes after 24 quarters, whereas, in the Netherlands it needs twelve quarters. On the other hand, a rise in $up$, panels (e) and (f), also reduces unemployment permanently in both countries, although in the Netherlands it has a positive short-run impact. Hence, only unemployment benefits have the expected long-run positive relation with unemployment. These results reinforce available evidence that not all LMI have a pernicious impact on unemployment (Nickell 1997, OECD 2006, Layard and Nickell 2011, Ch.7), and highlight the importance of considering specific shocks, a novelty of our paper, to inform policy design.

6.3 Long-term unemployment

We close out IRF-analysis by studying the impact of different supply and demand-shocks on long-term unemployment. It is sometimes argued that to correcting hysteresis caused by long-term unemployment requires supply-policies that increase incentives to work, reducing benefits or taxes, rather than positive demand shocks (de Koning et al. 2003, Tatsiramos and Ours 2014). To investigate this matter Fig. 13 assesses the response of
long-term unemployment to shocks on unemployment benefits, labour taxation, productivity, capital stock and interest rates.

Fig. 13. LMI and demand-shocks
Note: Solid lines denote IRF point estimates and dashed lines 95% CI.

A rise in unemployment benefits, panels (a) and (b), has no significant effect on $lu$ for the first 4-6 quarters in either country. However, after that, in the UK, long-term unemployment increases permanently by 22.5%, whereas, in the Netherlands $lu$ falls by 2%, although this reduction is only marginally significant. A rise in labour taxation, panels (c) and (d), also has country-specific effects. In the UK, the shock reduces long-term
unemployment permanently, whereas, in the Netherlands $lu$ stabilizes at a rate 10% above its pre-shock level. On the other hand, increases in productivity, panels (e) and (f), and capital stock, panels (g) and (h), have no significant effects on impact, but after 4-8 quarters these shocks reduce long-term unemployment permanently by 18-25%. Further, an increase in real long-term interest rates in the UK, panel (i), also has a negative effect on $lu$, whereas, in the Netherlands, panel (j), it has no significant impact.

Hence, Fig.13 suggests that labour market reforms that increase incentive to work can reduce long-term unemployment, although these need to be country-specific (reducing benefits in the UK and taxes in the Netherlands). Further, positive demand-shocks that encourage productivity and investment have similar results.

7. Summary

This paper investigated if there is hysteresis in the UK and the Netherlands and what demand variables can cause such phenomenon. We analysed this issue using a VAR-IRF model that extends available studies by testing for hysteresis through unemployment duration, productivity, capital stock and real long-term interest rates. Further, we also investigated what specific LMIs affect the NAIRU and the impact of supply and demand-shocks on long-term unemployment.

We find evidence of hysteresis in both countries, and that it happens through different channels. In the UK, it is due to the impact of productivity, capital stock and real long-term interest rates reduce unemployment permanently. In the Netherlands, hysteresis occurs through long-term unemployment, productivity and capital stock affect the NAIRU. Further, exploiting the wealth of estimates provided by our VAR-system, we find that productivity reduces the NAIRU because workers are sluggish to absorb improvements in productivity. The capital stock-NAIRU link seems to be the result of capital-scraping during downturns, and the impact of long-term unemployment in the Netherlands is due to the influence of unemployment duration on real wages.

On the other hand, we also find that the only LMI that increases unemployment permanently is unemployment benefits, although this effect is only marginally significant in the Netherlands. Further, we also find that long-term unemployment can be brought down by reducing unemployment benefits in the UK, and labour taxation in the Netherlands. However, positive demand-shocks of productivity or capital stock can have similar results.

Thus, our results contradict the belief that only LMI-reforms can reduce unemployment in the long-run. Instead our findings, suggest that policy makers have a choice between labour market reforms, which must be selective and country specific, and macroeconomic policies, which encourage productivity, investment and the reduction of long-term unemployment.
Appendix 1. Data

Table 1. Data description and sources

| Variables | Description | Source |
|-----------|-------------|--------|
| w−p | (log) real wage computed as w−p | [OC] |
| w | (log) average nominal wage: log (Employee Compensation / Total employment) | [2]/[1] |
| p | (log) Consumer Price Index (CPI) (Base 2005=1) for all items and the whole economy | [1] |
| y−1 | (log) real labour productivity: log (GDP / Total employment) | [OC] |
| GDP | Gross Domestic Product in nominal terms, in National Currency Units. | [2] |
| u | (log) unemployment rate (Labour Force Survey) | [2] |
| lu−1 | (log) long-term unemployment rate (Labour Force Survey): log ([Number of long-term unemployed workers] / [Number of unemployed workers]) * 100 | [1]/[2] |
| grr−1 | (log) linked Gross Replacement Rate calculated as the ratio between out-of-work benefits (under three family situations and three durations of unemployment) and in-work earnings (100% and 67% of manufacturing wages) times hundred. | [3] |
| t−1 | (log) linked Tax-wedge calculated as the ratio of taxation paid by workers over average labour costs, for a worker earning 100% of average wages under two family situations (single no children and married couple with one earner and two children): log [income tax + employees social security + employer's social security + cash transfers] / [gross earnings + employer social security] times hundred. | [4]/[1] |
| up (UK) | (log) strike activity, measured in number of days lost in labour dispute | [5] |
| up (Neth)−1 | (log) trade union density: log ([wage and salary earners that are trade union members] / [total number of wage and salary earners]) * 100 | [1] |
| k | (log) real capital stock for total economy (excluding housing services) in millions of local currency. | [2] |
| i−Ap | (log) real long-term interest rates: log (1 + (i−Ap * 100)) | [OC] |
| l10y | Long-term interest rates on government bonds | | |
| p−100 | (log) real price of imported raw materials computed as the ratio of Average Petroleum Spot index (in local currency) to CPI, weighted by share of imports to GDP: p−100 = v * log (P−100/CPI) | [OC] |
| v | Share imports to GDP (in percentage): (Imported goods and services/GDP) * 100 | [2] |
| P−100 | Average Petroleum Spot index in local currency. | [OC] |
| O10p | Average Petroleum Spot index of UK, Brent, Dubai & West Texas (Base 2005=1, in USD) | [6] |
| gbpusd | Exchange rate, USD per Great Britain Pound | [2] |
| eurusd | Exchange rate, USD per Euro | [2] |
| D95q2 | Netherlands’ dummy. D95q2=1 in 1995q2, D95q2=0 elsewhere | [OC] |
| T | Time-trend | [OC] |

Source: [OC] Own Calculation, [1] OECD.stat, [2] OECD Economic Outlook no. 90, [3] OECD Benefits and Wages indicators, [4] Correspondence with OECD’s Centre for Tax Policy and Administration, [5] ONS, [6] IMF.

Note: † Indicates that original annual data are transformed into quarterly data using linear interpolation. ‡ Indicates that original monthly data are transformed into quarterly data by considering the last month of the quarter observation as the quarterly value.

Appendix 2. Complementary results

Table 1. Lag order selection criteria

| Lag order | AIC | SBC | AIC | SBC |
|-----------|-----|-----|-----|-----|
| 1         | -46.72 | -43.81 | -54.25 | -51.12 |
| 2         | -50.76 | -45.84* | -57.44* | -52.30* |
| 3         | -50.80* | -43.87 | -57.00 | -49.85 |
| 4         | -50.37 | -41.43 | -57.05 | -47.88 |

Note: *Denotes criteria suggestion. Statistics obtained from estimating an unrestricted VAR-model for zt, with constant, time-trend, two lags of Δp−100, and the dummy D95q2 in the Netherlands. For both countries we use 110 observations, 1984q3-2011q4.
### Table 2. Long-run relationships

**United Kingdom, 1984q3-2011q4, 110 observations**

| $z_t^*$ | (1) | (2) | (3) | (4) | (5) |
|---------|-----|-----|-----|-----|-----|
| $(w-p)_t$ | 0.000 | -1.000 | 0.000 | -1.000 | 0.000 |
| $(y-l)_t$ | 0.000 | 1.000 | 0.000 | 1.000 | -1.013 |
| $u_t$ | -1.000 | 0.000 | -1.000 | 0.000 | -1.000 |
| $lu_t$ | 0.000 | 0.000 | 1.113 | 0.502 | 0.000 |
| $grr_t$ | 1.406 | 0.234 | 0.094 | -0.372 | 0.000 |
| $t_t^*$ | 2.526 | -0.114 | -0.590 | -1.511 | 1.680 |
| $mil_t$ | -0.018 | 0.049 | 0.480 | 0.003 | -0.129 |
| $k_t$ | 0.000 | 0.000 | 0.000 | 0.000 | 0.263 |
| $(i-\Delta p)_t$ | 0.000 | 0.000 | 0.000 | 0.000 | -0.000 |
| $T$ | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |

| $\chi^2_d(q-r^2)$ | $X^2_{(0.05)=5.2351}$ | $X^2_{(0.01)=19.742}$ | $X^2_{(0.00)=28.807}$ | $X^2_{(0.00)=43.392}$ | $X^2_{(0.00)=10.3075}$ |
|---------------|-----------------|-----------------|-----------------|-----------------|-----------------|
| $q$ | 14 | 12 | 11 | 12 | 14 |
| $\hat{\Pi}_d$ | 2869.6 | 2865.9 | 2881.4 | 2874.1 | 2890.6 |

**Netherlands, 1984q3-2011q4, 110 observations**

| $z_t^*$ | (1) | (2) | (3) | (4) | (5) |
|---------|-----|-----|-----|-----|-----|
| $(w-p)_t$ | 0.000 | -1.000 | 0.000 | -1.000 | 0.000 |
| $(y-l)_t$ | 0.000 | 1.000 | 0.000 | 1.000 | -1.013 |
| $u_t$ | -1.000 | 0.000 | -1.000 | 0.000 | -1.000 |
| $lu_t$ | 0.000 | 0.000 | 6.746 | 0.155 | 0.000 |
| $grr_t$ | 0.312 | 0.049 | -0.723 | -0.147 | -2.334 |
| $t_t^*$ | 3.191 | 0.263 | -0.201 | -0.296 | -2.304 |
| $tud_t$ | -0.516 | -0.070 | 12.59 | 0.279 | -0.580 |
| $k_t$ | 0.000 | 0.000 | 0.000 | 0.000 | -0.929 |
| $(i-\Delta p)_t$ | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| $T$ | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |

| $\chi^2_d(q-r^2)$ | $X^2_{(0.05)=56.647}$ | $X^2_{(0.01)=56.619}$ | $X^2_{(0.00)=44.440}$ | $X^2_{(0.00)=49.234}$ | $X^2_{(0.00)=9.532}$ |
|---------------|-----------------|-----------------|-----------------|-----------------|-----------------|
| $q$ | 14 | 12 | 11 | 12 | 11 |
| $\hat{\Pi}_d$ | 3248.4 | 3248.4 | 3254.5 | 3252.1 | 3272.0 |

Note: Asymptotic standard errors (SE) for $\beta$ coefficients are in brackets. * indicates significant at 5% and * at 10%. (NC) indicates failure to produce converging results when computing SE. $\hat{\Pi}_d$ is the maximum value of the log-likelihood function under $q$ over-identifying restrictions. $r^2$ is the number of just-identified. UK, $\hat{\Pi}_d = 2895.8$. Netherlands, $\hat{\Pi}_d = 3276.7$. $\chi^2_d(q-r^2)$ is chi-square statistic for the LR test, p-values for this test are in square brackets.
Table 3. Unemployment short-run dynamics

|               | UK 1984q3-2011q4 | Netherlands 1984q3-2011q4 |
|---------------|-------------------|--------------------------|
|               | $\Delta u_t$     | $\Delta u_t$              |
| $\xi_{zt-1}$  | -0.022 [0.438]    | -0.031$^{\dagger}$ [0.009] |
| $\xi_{zt-1}$  | 0.211$^{\dagger}$ [0.004] | 0.482 [0.156] |
| $c_0$         | -0.429$^{\dagger}$ [0.008] | 5.274$^{\dagger}$ [0.037] |
| $\Delta(w_{t-1} - p_{t-1})$ | -0.431 [0.268] | 0.296 [0.472] |
| $\Delta(y_{t-1} - l_{t-1})$ | -0.487 [0.171] | -0.603$^{\dagger}$ [0.049] |
| $\Delta u_{t-1}$ | 0.395$^{\dagger}$ [0.000] | 0.079$^{\dagger}$ [0.000] |
| $\Delta u_{t-1}$ | -0.236$^{\dagger}$ [0.025] | -0.115$^{\!*}$ [0.068] |
| $\Delta gr_{t-1}$ | 1.369$^{\dagger}$ [0.008] | 0.076 [0.701] |
| $\Delta t_{t-1}^2$ | -0.045 [0.906] | -0.047 [0.797] |
| $\Delta u_{t-1}$ | 0.002 [0.460] | -0.151 [0.632] |
| $\Delta k_{t-1}$ | -2.466$^{\!*}$ [0.071] | -3.902$^{\dagger}$ [0.025] |
| $\Delta (l_{t-1} - \Delta p_{t-1})$ | 0.032$^{\dagger}$ [0.009] | 0.009 [0.424] |
| $\Delta p_{t-1}^{2m}$ | -0.000 [0.604] | -0.000 [0.250] |
| $\Delta p_{t-1}^{2m}$ | -0.000 [0.342] | -0.000 [0.572] |
| $D^95q2$      | n/a               | 0.121$^{\dagger}$ [0.000] |

Observations 110 110

Adj. $R^2$ 0.522 0.768

$X_{EC}^2(4)$ 8.452$^{*}$ 3.817

$X_{FF}^2(1)$ 2.754$^{*}$ 0.048

$X_{Norm}^2(2)$ 4.354 12.45$^{\!*}$

$X_{Het}^2(1)$ 0.001 0.108

Note: $p$-values for $t$-tests and diagnostic tests in square brackets. $^{\dagger}$ indicates significant at 5% and $^{*}$ at 10%. Adj. $R^2$=Adjusted R-square. $X_{EC}^2$, $X_{FF}^2$, $X_{Norm}^2$, $X_{Het}^2$ are chi-square statistics for Serial correlation (SC), Functional Form (FF), Residuals Normality (NORM) and Heteroscedasticity (HET) tests, respectively.

For the UK:

$\xi_{zt} = u - 0.832u + 0.147k$

$\xi_{zt} = (w - p) - (y - f) - 0.253lu + 1.375\sigma + 0.262u + 0.135(i - \Delta p)$

For the Netherlands:

$\xi_{zt} = u - 2.054u + 5.226\sigma + 2.928l + 1.594(i - \Delta p)$

$\xi_{zt} = (w - p) - 0.300(y - f) + 0.101gr + 0.273up - 0.484k + 0.054(i - \Delta p) + 0.0057$

Fig. 1. Eigenvalues of the companion matrix
## Appendix 3. Unit root and stationarity tests

### Table 1 ADG-GLS, UK

| Variable | ADF-GLS(1) | ADF-GLS(2) | ADF-GLS(3) | ADF-GLS(4) | ADF-GLS(5) |
|----------|------------|------------|------------|------------|------------|
| (i) First differences |           |            |            |            |            |
| ∆(w_t - p_t) | -5.638     | -4.770     | -3.329     | -3.732     | -2.987     |
| ∆(y_t - l_t) | -6.389     | -5.302     | -3.464     | -3.956     | -3.062     |
| ∆u_t | -2.814     | -3.032     | -3.372     | -3.456     | -3.036     |
| ∆l_t | -3.400     | -3.521     | -3.677     | -2.664     | -2.693     |
| ∆r_t | -2.569     | -2.633     | -2.733     | -2.941     | -2.943     |
| ∆t^w | -2.943     | -3.097     | -3.273     | -2.919     | -3.093     |
| ∆m_t | -4.325     | -2.970     | -2.493     | -1.868     | -1.356     |
| ∆k_t | -0.819     | -0.956     | -1.081     | -1.162     | -1.229     |
| ∆(l_t - Δp_t) | -6.559     | -4.882     | -2.795     | -2.456     | -1.798     |
| ∆p^m | -9.348     | -6.197     | -5.285     | -5.659     | -4.707     |
| (ii) Levels |           |            |            |            |            |
| w_t - p_t | -0.223     | -0.746     | -0.806     | -1.346     | -1.008     |
| y_t - l_t | -0.213     | -0.611     | -0.641     | -1.258     | -0.845     |
| u_t | -1.212     | -1.907     | -1.620     | -1.204     | -0.987     |
| l_t | -2.560     | -2.407     | -2.223     | -2.008     | -2.627     |
| g_t | -2.243     | -2.191     | -2.075     | -1.921     | -1.683     |
| t^w | -3.396     | -3.294     | -3.117     | -2.906     | -3.211     |
| m_t | -9.676     | -4.497     | -4.203     | -3.405     | -2.984     |
| k_t | -2.381     | -2.724     | -2.498     | -2.310     | -2.210     |
| i_t - Δp_t | -3.691     | -3.236     | -2.309     | -2.596     | -2.099     |
| p^m | -9.348     | -6.197     | -5.285     | -5.659     | -4.707     |

Note: ADF-GLS(p) is Elliott et al. (1996) GLS augmented Dickey-Fuller unit root statistic for p-lags. (i) 1st difference, ADF-GLS statistics computed using p-lagged 1st differences of dependent variable and an intercept. (ii) Level, ADF-GLS statistics computed using p-lags of dependent variable, intercept and time-trend, except for ρ where no time-trend is used. 5% critical value without trend=-1.950, with trend=-3.017. 1983q4-2011q4, 113 observations.

### Table 2 KPSS, UK

| Variable | KPSS(0) | KPSS(2) | KPSS(4) | KPSS(6) | KPSS(8) | KPSS(10) | KPSS(12) |
|----------|---------|---------|---------|---------|---------|----------|----------|
| (i) First differences |           |         |         |         |         |          |          |
| ∆(w_t - p_t) | 0.496 | 0.473 | 0.397 | 0.367 | 0.329 | 0.293 | 0.266 |
| ∆(y_t - l_t) | 0.396 | 0.465 | 0.412 | 0.393 | 0.365 | 0.337 | 0.317 |
| ∆u_t | 1.150 | 0.505 | 0.361 | 0.310 | 0.287 | 0.273 | 0.261 |
| ∆l_t | 0.752 | 0.297 | 0.208 | 0.172 | 0.153 | 0.142 | 0.136 |
| ∆r_t | 1.590 | 0.601 | 0.402 | 0.323 | 0.287 | 0.269 | 0.259 |
| ∆t^w | 0.370 | 0.142 | 0.097 | 0.079 | 0.071 | 0.067 | 0.066 |
| ∆m_t | 0.033 | 0.058 | 0.092 | 0.115 | 0.141 | 0.149 | 0.171 |
| ∆k_t | 2.080 | 0.734 | 0.462 | 0.347 | 0.284 | 0.246 | 0.222 |
| ∆(l_t - Δp_t) | 0.026 | 0.072 | 0.101 | 0.132 | 0.162 | 0.178 | 0.185 |
| ∆p^m | 0.243 | 0.301 | 0.325 | 0.405 | 0.426 | 0.462 | 0.460 |
| (ii) Levels |         |         |         |         |         |          |          |
| w_t - p_t | 0.888 | 0.320 | 0.204 | 0.156 | 0.129 | 0.113 | 0.102 |
| y_t - l_t | 1.140 | 0.412 | 0.264 | 0.201 | 0.168 | 0.147 | 0.134 |
| u_t | 1.260 | 0.438 | 0.275 | 0.207 | 0.170 | 0.149 | 0.135 |
| l_t | 0.748 | 0.260 | 0.164 | 0.125 | 0.104 | 0.092 | 0.085 |
| g_t | 0.660 | 0.230 | 0.146 | 0.112 | 0.094 | 0.084 | 0.079 |
| t^w | 0.566 | 0.193 | 0.121 | 0.091 | 0.076 | 0.067 | 0.062 |
| m_t | 0.750 | 0.412 | 0.330 | 0.282 | 0.246 | 0.222 | 0.207 |
| k_t | 1.200 | 0.393 | 0.247 | 0.185 | 0.152 | 0.131 | 0.118 |
| i_t - Δp_t | 0.322 | 0.231 | 0.178 | 0.154 | 0.139 | 0.128 | 0.121 |
| p^m | 3.560 | 1.290 | 0.821 | 0.620 | 0.507 | 0.434 | 0.383 |

Note: KPSS(l) is Kwiatkowski et al. (1992) stationarity-test with Bartlett window=l. 113 observations, 1983q4-2011q4. (i) 1st difference, KPSS statistics computed using p-lagged 1st differences of dependent variable and an intercept. (ii) Level, KPSS statistics computed using p-lags of dependent variable, intercept and time-trend, except for p^m where no time-trend is used. 5% critical value without trend=0.463, with trend=0.146. 1983q4-2011q4, 113 observations.
Table 3 ADG-GLS, Netherlands

| Variable | ADF-GLS(1) | ADF-GLS(2) | ADF-GLS(3) | ADF-GLS(4) | ADF-GLS(5) |
|----------|------------|------------|------------|------------|------------|
| $\Delta (w_t - p_t)$ | -2.388 | -1.531 | -1.029 | -0.676 | -0.604 |
| $\Delta (y_t - l_t)$ | -7.587 | -5.712 | -4.907 | -4.559 | -4.670 |
| $\Delta u_t$ | -2.955 | -3.076 | -3.298 | -2.914 | -2.997 |
| $\Delta u_t$ | -3.759 | -3.804 | -3.938 | -2.979 | -2.939 |
| $\Delta grr_t$ | -1.731 | -1.740 | -1.743 | -1.757 | -1.777 |
| $\Delta t^*_n$ | -4.229 | -4.622 | -5.215 | -4.570 | -4.869 |
| $\Delta u_t$ | -2.862 | -3.033 | -3.118 | -1.953 | -1.939 |
| $\Delta k_t$ | -1.126 | -1.555 | -1.698 | -1.549 | -1.521 |
| $\Delta (l_t - \Delta p_t)$ | -5.534 | -3.422 | -2.208 | -1.564 | -1.274 |
| $\Delta p_t^m$ | -9.205 | -5.925 | -5.196 | -5.712 | -4.880 |

Note: ADF-GLS($p$) is Elliott et al. (1996) GLS augmented Dickey-Fuller unit root statistic for $p$-lags. (i) 1st difference, ADF-GLS statistics computed using $p$-lagged 1st differences of dependent variable and an intercept. (ii) Level, ADF-GLS statistics computed using $p$-lags of dependent variable, intercept and time-trend, except for $p_t^m$ where no time-trend is used. 5% critical value without trend=-1.950, with trend=-3.017. 1983q4-2011q4, 113 observations.

Table 4 KPSS, Netherlands

| Variable | KPSS(0) | KPSS(2) | KPSS(4) | KPSS(6) | KPSS(8) | KPSS(10) | KPSS(12) |
|----------|---------|---------|---------|---------|---------|---------|---------|
| $\Delta (w_t - p_t)$ | 0.160 | 0.220 | 0.220 | 0.196 | 0.190 | 0.189 | 0.194 |
| $\Delta (y_t - l_t)$ | 0.081 | 0.106 | 0.106 | 0.113 | 0.120 | 0.119 | 0.119 |
| $\Delta u_t$ | 0.754 | 0.294 | 0.202 | 0.166 | 0.149 | 0.142 | 0.141 |
| $\Delta u_t$ | 0.240 | 0.105 | 0.081 | 0.074 | 0.072 | 0.073 | 0.077 |
| $\Delta grr_t$ | 1.890 | 0.712 | 0.475 | 0.379 | 0.333 | 0.309 | 0.295 |
| $\Delta t^*_n$ | 0.171 | 0.074 | 0.058 | 0.057 | 0.065 | 0.080 | 0.101 |
| $\Delta u_t$ | 0.553 | 0.224 | 0.163 | 0.138 | 0.124 | 0.114 | 0.108 |
| $\Delta k_t$ | 1.340 | 0.470 | 0.299 | 0.230 | 0.194 | 0.174 | 0.162 |
| $\Delta (l_t - \Delta p_t)$ | 0.009 | 0.029 | 0.045 | 0.060 | 0.075 | 0.078 | 0.087 |
| $\Delta p_t^m$ | 0.221 | 0.257 | 0.266 | 0.326 | 0.350 | 0.387 | 0.400 |

Note: KPSS($l$) is Kwiatkowski et al. (1992) stationarity-test with Bartlett window=$l$ (i) 1st difference, KPSS-statistics computed with and intercept and $l$ lagged truncation parameter. (ii) Level, KPSS-statistics computed with and intercept, time-trend and $l$ lagged truncation parameter, except in the case of $p_t^m$ where no time-trend is considered. 5% critical value without trend=-0.463, and with trend=0.146. 1983q4-2011q4, 113 observations.
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