Examining the Relationship Between Prices in the Netherlands and the Euro-area

Theory, Evidence and Forecast

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Abstract

Understanding the nature of the relationship between prices in two members of a currency union, which are not members of a common fiscal union, is of utmost importance for successful macroeconomic policy. This paper examines such a relationship in the context of the Euro-area. The focus is on the relationship between inflation and price levels in the Netherlands and the rest of the Euro-area. This work contributes to the current literature in several ways. First, it develops an alternative empirical strategy for estimating inflation based on the aforementioned relationship. Second, this work provides quite extensive empirical tests of this relationship. Third, the possibility of using these models for forecasting is examined as well.

Key words: inflation, currency union, Euro-area, ARDL, error correction.
JEL classification: E31; E37; F36.

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Contents

1 Introduction

On January 1, 1999, the Netherlands formally gave up her independent monetary policy, following her ascension to the Euro-area. Three years later, on January 1, 2002, Netherlands introduced Euro currency into circulation and abandoned the guilder just 28 days later. However, these events did not make the national Dutch inflation rate and price level any less relevant for the public policy. On the contrary, the local inflation rate and price level became more important for public policy than ever before.

The reasons for this are the economically important and persistent inflation rate and price level differentials between members of the Euro-area (Beck et al., 2009). These are important because inflation rate and price levels have a large impact on the macroeconomic environment, but the Euro-area shares a single monetary policy. In such an environment, understanding the relationship between the national rate of inflation and price level, and those of the rest of the Euro-area become crucially important. The large and asymmetric macroeconomic shocks which are currently impairing economic activity in the Euro-area only add to the importance of this relationship for policy makers.

Still, despite of such relevance, surprisingly little research was done on this topic before. The purpose of this paper is to fill this gap, by exploring and estimating the relationship between inflation rates and price levels amongst members of the currency area. Furthermore, this research will be applied to the Netherlands, and will also include forecasting models for the Dutch inflation.

Prior to the Great Recession, research mainly focused on optimal policy rules in currency areas (see Ca'Zorzi et al., 2005, Ferrero, 2009, Lombardo, 2006). Afterwards researchers were preoccupied more with the persistence of inflation differentials among various Euro-zone countries (See Altissimo et al., 2011, Angeloni and Ehrmann, 2007, Beck et al., 2009), than with the relationship between national and 'aggregate' inflation per se.

Nevertheless, this is not due to a lack of foundation to build upon. As a matter of fact, prior literature already provides a solid base for this paper. For example, the two-sector two-country general equilibrium model of Ca'Zorzi et al. (2005) already indirectly postulates a theoretical relationship between the price levels of two common currency area members. Moreover, the empirical research on inflation differentials of Altissimo et al. (2011), Angeloni and Ehrmann (2007), Beck et al. (2009) provides a good starting point for this research.

The rest of the paper is organised as follows. The second section will present the theoretical model of the relationship between price levels in the Netherlands and the rest of the Euro-area, based on Ca'Zorzi et al. (2005). The third section will show the identification strategy for the estimation of this relationship. Section four will display the results of the empirical estimation. Section five will develop and evaluate forecasting models based on previous sections. The sixth section will discuss possible extensions and the seventh section will conclude.

Moreover, some of these sections can be (partially) skipped without a loss of cohesion. For example, some readers may find it convenient to skip subsection 3.1 which provides details on data selection. Furthermore, readers not interested in econometric details may skip most of section 4 and can go directly to the overview of results in subsection 4.5. Those not interested in the forecast can omit the section 5.

2 Theoretical Background

This section will provide a theoretical background for the estimation of the relationship between inflation in the Netherlands and the rest of the Euro-area. Moreover, this part will borrow heavily from the two-sector, two-country general equilibrium model of Ca'Zorzi et al. (2005). Their model will serve well the purposes of this paper because it is relatively simple and it was developed to describe the Euro-area. Furthermore, in this section, we will follow their model closely, but we will apply it to the Netherlands and extend it to hold intertemporally as well.

Here, we will assume that the Netherlands (denoted by NL) is a small open economy that has joined a currency union with a large economy, the Euro-area (denoted by EA). Both of them are producing traded goods (T) and non-traded goods (N).

In contrast to Ca'Zorzi et al. (2005) this section provides a little bit more micro foundations. We will start by deriving the indirect money metric utility function, following the Varian (1992), McKenzie (1957) and Samuelson (1974). The starting point for this will be the Cobb-Douglas utility function of a representative individual,

$$U_i\left(Y_{iT}, Y_{iN}\right) = Y_{iT}^{\gamma_i} Y_{iN}^{1-\gamma_i} \tag{1}$$

which will be maximized subject to the following budget constraint:

$$M_i = P_{iT} Y_{iT} + P_{iN} Y_{iN} \tag{2}$$

where Y_{ij} and P_{ij} are the traded and non-traded goods and their respective prices, ($j = \{T, N\}$) in the Netherlands and the rest of the Euro-area[1] ($i = \{NL, EA\}$). Because any monotonic transformation of a utility function still represent the same preferences we work with the log transformation of the utility function,

$$U_i\left(Y_{iT}, Y_{iN}\right) = \gamma_i \ln Y_{iT} + (1-\gamma_i) \ln Y_{iN} \tag{3}$$

The Marshallian demands can be obtained by maximizing this utility function subject to the budget constraint. The first order conditions are given by

$$\frac{\gamma_i}{Y_{iT}} - \lambda P_{iT} = 0 \tag{4}$$

and

[1] Note that unless stated otherwise we will be concerned here with the Euro-area excluding the Netherlands itself.

$$\frac{1-\gamma_i}{Y_{iN}} - \lambda P_{iN} = 0 \tag{5}$$

and the equation 2 that describes the budget constraint. These expressions can be combined to get

$$\frac{\gamma_i}{P_{iT}Y_{iT}} = \frac{1-\gamma_i}{P_{iN}Y_{iN}} \tag{6}$$

Cross-multiplying the first order conditions produces

$$\gamma_i P_{iN} Y_{iN} = P_{iT}Y_{iT} - \gamma_i P_{iT} Y_{iT} \tag{7}$$

Rearranging and substituting the budget constraint gives

$$\gamma_i M = P_{iT} Y_{iT} \tag{8}$$

This expression can be further rearranged in a following way

$$Y_{iT}(P_{iT}, P_{iN}, M_i) = \frac{\gamma_i M_i}{P_{iT}} \text{ or } Y_{iN}(P_{iT}, P_{iN}, M_i) = \frac{1-\gamma_i M_i}{P_{iN}} \tag{9}$$

Substituting back to the objective function gives the indirect utility function

$$v_i(P_{iT}, P_{iN}, M_i) = \gamma_i^{\gamma_i}(1-\gamma_i)^{1-\gamma_i}\frac{M_i}{P_{iT}^{\gamma_i}P_{iN}^{1-\gamma_i}} \tag{10}$$

From this it follows that the true price deflator is $P = P_{iT}^{\gamma_i}P_{iN}^{1-\gamma_i}$, since this reflect the shape of utility function. We can get rid of the constant as any monotonic transformation represents the same preferences. Finally, the money metric utility function can be derived by substitution

$$M_i(P_{ij}, Y_{ij}) = P_{iT}^{\gamma_i}P_{iN}^{1-\gamma_i}U_i(Y_{iT}, Y_{iN}) = P_{iT}^{\gamma_i}P_{iN}^{1-\gamma_i}Y_{iT}^{\gamma_i}Y_{iN}^{1-\gamma_i} \tag{11}$$

Taking logs of the equation 11 produces more useful expression:

$$m = \gamma_i p_{iT} + (1-\gamma_i)p_{iN} + \gamma_i y_{iT} + (1-\gamma_i)y_{iN} \tag{12}$$

where the small case letters denote logs of variables. Moreover, from the macroeconomic perspective, the budget can be interpreted as the aggregate product at aggregate prices ($m = p_i + y_i$), because one person's expenditure is another person's income.

Since income can be spent either on traded or non-traded goods, assuming homothetic preferences, we can express the aggregate income (in logarithms) as the sum of the income spent on traded (T) and non-traded (N) goods:

$$y_i = \gamma_i y_{iT} + (1 - \gamma_i) y_{iN} \tag{13}$$

Where γ_i gives the respective share of income (y) spent on either of those.

The consumer price index [CPI] can be expressed as a weighted average of prices in the traded (p_{iT}) and non-traded (p_{iN}) sector. The individual weights in the CPI are given by the share of income spent on tradables and non-tradables (the lower-case letters stand for the natural logs of variables). In logarithmic form the price deflator in the Netherlands and the rest of the Euro-area is

$$p_i = \gamma_i p_{iT} + (1 - \gamma_i) p_{iN} \tag{14}$$

Moreover, in this setting the sectoral demand for goods will be given by the relative prices, where $j = \{TN\}$:

$$y_{ij} - y_i = -(p_{ij} - p_i) \tag{15}$$

Notice that this specification simply states that the relative differences in demand for tradable and non-tradable output will be given by the differences in the relative prices of tradables and non-tradables. To show this we start by rearranging the equation 15:

$$p_i + y_i = p_{ij} + y_{ij} \tag{16}$$

Next we can substitute the equation 13 and 14 for p_i and y_i:

$$\gamma_i p_{iT} + (1 - \gamma_i) p_{iN} + \gamma_i y_{iT} + (1 - \gamma_i) y_{iN} = p_{ij} + y_{ij} \tag{17}$$

Regardless of choosing $j = N$ or T rearranging the equation 17 gives:

$$y_{iT} - y_{iN} = -(p_{iT} - p_{iN}) \tag{18}$$

After describing the demand side, we can turn our attention to the supply side of the Ca'Zorzi et al. (2005) model. Using the Cobb-Douglas production function and normalizing the capital stock to 1, we can get sectoral output[2]:

[2]Remember that everything is expressed in natural logs, so here we will work with a log-linearized version of $A_i K_i^b L_i^{1-b}$, where K is normalized to 1.

$$y_{ij} = a_{ij} + (1 - b_{ij})l_i \tag{19}$$

Where a_{ij} is the log of productivity. $(1 - b_{ij})$ can be interpreted as the factor input elasticity of labour. The log of labour supply is denoted as l_i. Aggregating over all sectors gives:

$$y_i = \gamma_i a_{iT} + (1 - \gamma_i)a_{iN} + \gamma_i(1 - b_{iT})l_{iT} + (1 - \gamma_i)(1 - b_{iT})l_{iN} \tag{20}$$

The sectoral demands on individual labour markets can be expressed by equating producers real wages to the marginal product of labour[3] (as these have to equal in the equilibrium):

$$l_{ij} = p_{ij} + y_{ij} - w_i \tag{21}$$

Combining equation 15 and 21 leads to:

$$p_{iN} - p_{iT} = (y_{iT} - l_{iT}) - (y_{iN} - l_{iN}) \tag{22}$$

Because relative prices adjust to prevent a change in the sectoral allocation of employment, we can combine equation 19 with the previous equation to get:

$$p_{iN} - p_{iT} = a_{iT} - a_{iN} + (b_{iN} - b_{iT})l_i \tag{23}$$

Allowing for small deviations (ϵ) from the purchasing power parity [PPP] in the Dutch tradable sector, it is possible to describe the real exchange rate q_{iT} as

$$q_{NL,T} = p_{EA,T} + \epsilon - p_{NL,T} \tag{24}$$

Using equation 23, equation 14 can be rewritten to show that:

$$p_{NL} = p_{EA,T} + \epsilon + \theta_{NL} \tag{25}$$

and also that:

$$p_{EA} = p_{EA,T} + \theta_{EA} \tag{26}$$

Where θ_i is defined by equation 27 below, and introduced to make the expression more compact. The θ_i represents the structural parameters of the economy. It

[3]We get to this expression by taking a log of both sides of $Y_{ij} = \frac{L_{ij}W_i}{P_{ij}}$ and solving for a log of labour supply. This relationship must hold in competitive equilibrium where profits are 0, using the following profit function $\mathcal{P}$ of the representative firm; $\mathcal{P}_i(L_i) = P_{ij}Y_{ij}(L_{ij}) - W_iL_{ij}$.

consists of the productivity differentials $a_{iT} - a_{iN}$, labor supply $(b_{iN} - b_{iT})l_i$, and it is weighted by the share of income spent on non-traded goods.

$$\theta_i = (1 - \gamma_i)[a_{iT} - a_{iN} + (b_{iN} - b_{iT})l_i] - q_{iT} \tag{27}$$

Now solving equation 26 in terms of $p_{EA,T}$ and substituting the solution into equation 25 we get:

$$p_{NL} = p_{EA} - \theta_{EA} + \theta_{NL} + \epsilon \tag{28}$$

This last relationship describes the relationship between price levels in both the Netherlands and the rest of the Euro-area directly as a function of the overall price level, and their respective structural factors[4]. To understand the intuition of the model we can compare it to the classical Balasa-Samuelson effect. The Balassa-Samuelson (1964) effect postulates that countries with high productivity growth also experience high wage growth, leading to higher real exchange rates. Thus an increase in wages in the tradable goods sector of an economy will also lead to higher wages in the non-tradable sector of the economy. The accompanying increase in inflation makes inflation rates higher in faster-growing economies.

Our result tells virtually the same story, in the absence of nominal exchange rate adjustment, price levels between the Netherlands and rest of the Euro-area need to adjust to prevent arbitrage between the two. Consequently, the price level differentials between the Netherlands and the rest of the Euro-area can be viewed as a real exchange rate. Thus, the persistent difference between the price levels in the two areas can be seen as a reflection of the Balasa-Samuelson theorem, as well as idiosyncratic shocks (Beck et al., 2009).

If price levels would be equal everywhere in the Euro-area, every small idiosyncratic shock to the marginal productivity would require labour or capital crossing borders for markets to equilibrate. However, such erratic movements of labour and capital can be prevented as long as price levels are allowed to adjust instead of the supply of inputs[5].

[4] Remember that $q_{iT} = 0$ by construction. This holds because under the law of one price log of nominal exchange s equals the difference of logs of price levels p_{NL}, p_{EA}, that is $s = p_{NL} - p_{EA}$. The real exchange rate is by definition $q \equiv s - p_{NL} + p_{EA}$ and thus under the law of one price $q \equiv 0$.

[5] Of course, in practice, real exchange rate might be hard to adjust due to market imperfections (such as price and wage stickiness). These could be in principle introduced into the model via equation 21. Because of this, capital and labour still move across the borders, to help markets equilibrate, but in less erratic fashion. However, such analysis is not necessary for the purposes of this paper, and thus it will not be pursued here.

Again, this is no different from the notion of real exchange rate. We can also see this more explicitly, since the exchange rate, based on the notion of PPP, is by definition the difference between the natural logs of the price level ($q \equiv p_{NL} - p_{EA}$). By subtracting the price level from both sides of equation 28, we can get the expression of the exchange rate as a function of structural factors (θ_i) and the idiosyncratic shocks (ϵ).

$$p_{NL} - p_{EA} = \theta_{NL} - \theta_{EA} + \epsilon \tag{29}$$

In fact, the literature preoccupied with the estimation of inflation and price level differentials is, loosely speaking, based on the equation 29[6]. However, this research is interested more in the direct relationship between the price levels. Thus, hereafter, we will work with the relationship described by equation 28.

[6] Although, this might not always be obvious at first sight.

3 Empirical Strategy

In practice, estimating the equation 28 directly would likely lead to spurious regression. The reason for this is that price indices and structural parameters as well are generally non-stationary[7]. Because of this (co)variances will be ill-defined leading to spurious regression (see Verbeek, 2008). Nevertheless, this particular problem can be fixed by taking the first differences of equation 28. To do this, we need to further assume that equation 28 holds for every period. Afterwards, we can simply subtract the first lag of equation 28 from itself. This gives the expression for the relationship between inflation rates in the Netherlands and the rest of the Euro-area.

$$\pi_{NL,t} = \pi_{EA,t} + \vartheta_{NL,t} - \vartheta_{EA,t} + \varepsilon_t \tag{30}$$

where $\pi_{i,t} = \Delta p_{i,t}$, $\vartheta_{i,t} = \Delta\theta_{i,t}$; $i = \{NL, EA\}$ and $\varepsilon_t = \Delta\epsilon_t$

However, equation 28 is still useful. We can think of it as describing the long run relationship between the two price levels. To get better estimates of coefficients from the equation 30, we can add an error correction term. This can be done by solving the equation 28 for the error (ϵ), and by adding its lag to the estimation of 30.

This approach has several further advantages. To begin with, adding such term would help to capture the long-term dynamics as well as provide further information on the rate of adjustment of the inflation toward the long run equilibrium. Moreover, estimating the error correction mechanism will also provide us with the estimates of the long-run relationship between the price levels.

We will also add lags of inflation rates to allow our model to capture short run dynamics in the inflation rates. Another addition are control variables represented by the vector z. These include mainly the dummies to account for the value added tax (VAT) increases in the Netherlands.

$$\pi_{NL,t} = \alpha + \sum_i \beta_i \pi_{NL,t-1-i} + \sum_i \Lambda_i \pi_{EA,t-i} + \upsilon\vartheta_{NL,t} - \nu\vartheta_{EA,t} - \mu\epsilon_{t-1} + \delta_i z_t + \varepsilon_t \tag{31}$$

However, the equation 31 is very hard to estimate directly because some of the components of ϑ_i are hard to observe. These and other issues pertaining to the data will be discussed in the following subsection.

[7]Which is in this particular case also confirmed by wide range of unit root tests on price indices.

3.1 Data Selection

This subsection provides the justification for the selection of the data for this research. The first part (3.1.1) will be dedicated to the data on inflation rates and price levels in the Netherlands and the rest of the Euro-area. The second part (3.1.2) will focus on the data selection for structural parameters. The third part (3.1.3) deals with the control dummies.

3.1.1 Inflation in the Netherlands and the Euro-area

The inflation is measured using the harmonized consumer price index (HICP), which is published by Eurostat, and it is the most commonly used index. All goods and services are included. With having monthly data on the HICP we can measure inflation in multiple ways, with two most common ways being the annualized monthly rate and the year-on-year rate.

The annualized monthly rate is defined as $[\left(\frac{HICP_t}{HICP_{t-1}}\right)^{12} - 1]$ or after taking logs as $12[ln(HICP_t) - ln(HICP_{t-1})]$. In addition the series are seasonally adjusted using the TRAMO/SEATS procedure. However, the downside of this measure is that the seasonal adjustment may also discard genuine variation in the data.

In contrast, the year-on-year rate, defined as $\left(\frac{HICP_t}{HICP_{T-12}} - 1\right)$ or after log transformation as $[ln(HICP_t) - ln(CPI_{t-12})]$, is a convenient way of seasonally adjusting the series (Carnot et al., 2011), as it does not discard any variation from the original data. The downside of this method is that it compares data across the year and thus we have to discard a year worth of observations from our dataset.

The monthly data for HICP of the Netherlands and the Euro-area were taken from the Eurostat's HICP database[8]. Moreover, the inflation in the Euro-area was adjusted to exclude the Netherlands. This was done using the weights published by the Eurostat[9], in a way that share of all the other Euro-area members was preserved. Unfortunately, the adjusted dataset ranges only from January 1996 to June 2016. Not only is this very short time period, but it also encompasses many structural changes.

Officially the Dutch guilder, as well as other currencies of the Euro-area founding members, ceased to exist on January 1, 1999[10]. However, until January 2002, Euro existed only on 'the books' as a physical Euro currency was introduced afterwards, with the guilder being finally relegated to the dustbin of history on January 28, 2002.

[8] This database is listed under code: prc_hicp_mv12r.

[9] This database is listed under code: prc_hicp_cow.

[10] Nevertheless, the Euro existed in a 'book form' already since the January 1, 1996.

During the transitory period the Netherlands was on fixed exchange which was set at €1 = 2.20371 NLG.

This introduces potential problems, as this seemingly innocent change can influence the inflation rate. For example, many observers were suspicious at the time about businesses using the currency switch to increase prices across the board (De Nederlandsche Bank, 2002). Furthermore, the Netherlands also increased its value added tax (VAT) in January 2001 from 17.5% to 19%, which was another significant structural change during the transition period of 1999-2002[11]. It turns out that these had a considerable impact on the Dutch inflation rate. This can be seen from the figure 1 below, which plots the year-on-year change in the inflation rate in the Netherlands and the rest of the Euro-area.

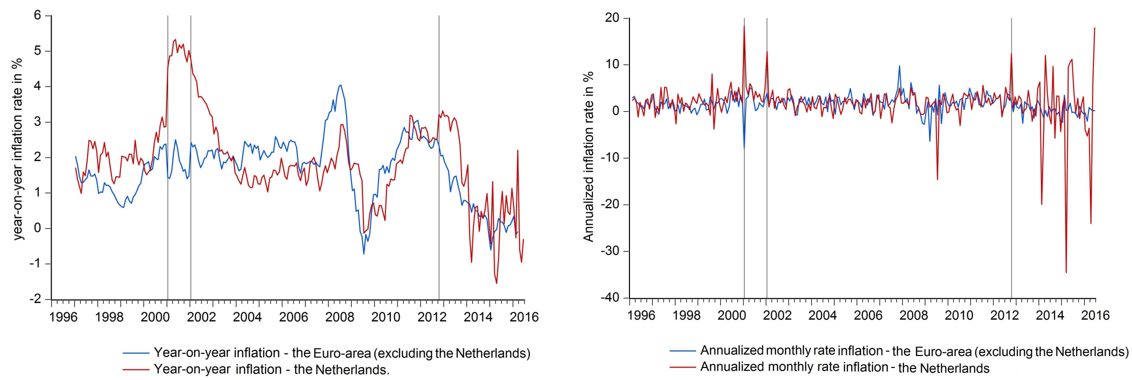


Figure 1: Plots of the year-on-year and annualized monthly rate of inflation in the Netherlands and the rest of the Euro-area.

The first vertical line represents the first VAT hike on January 1, 2001, from 17.5% to 19% (as well as the increase in the eco-tax). The second line denotes the introduction of the Euro as a physical currency. The third line shows the second VAT hike on October 1, 2012, from 19% to 21%. The figure 1 clearly shows huge, almost vertical spike in the prices in the year following the 2001 VAT hike, which is consistent with the findings of Mellens et al. (2014). Additionally, the figure shows that precisely when inflation started to fall during several consecutive months, there was another small visible spike around the time the Euro was physically introduced, which slowed down the decline in the price level throughout the 2002. Furthermore,

[11] Furthermore, there were also increases in other taxes in this period, such as in the eco-tax, which might have also contributed to the spike in inflation.

the inference is also made harder by the fact that the Great Recession spans a several years of the database.

Hikes in multiple tax rates followed by the introduction of the Euro in quick succession create several challenges during the data analysis. Perron (1989) argues that in the presence of structural break(s) unit root tests are biased into finding a unit root process even in stationary series. The introduction of Euro and tax hikes are likely to create structural breaks. The Chow test also indicates this by rejecting the null hypothesis that there was no structural break during that period.

However, while there are unit root tests that allow for multiple structural breaks, these have generally smaller power, despite being more consistent (Glynn et al., 2007). This can cause serious problem in an already very limited sample. Moreover, structural beaks can also lead to spurious unit roots in cointegration tests according to Beyer et al. (2009). Given these issues, it might be desirable to exclude all the years that preceded the physical adoption of the Euro from the sample.

The table 34 and 35 show a summary of a number of unit root tests for inflation rate in the Netherlands. The reason for calculating multiple unit root tests is that different information criteria selected a widely different number of lags, and unit root tests are sensitive to lag selection. The maximum number of lags for the criteria to consider was set to 36, based on the sample size.

Moreover, because the ADF test can have low power in small samples (Verbeek, 2008), KPSS tests were also calculated to control for this. Popular kernel specifications are the Bartlett weights and the quadratic spectral kernel. Hence, both are used for a robustness check. Serial correlation is corrected with Newey-West and Andrews correction respectively[12].

In the full sample ADF tests generally cannot reject the null hypothesis of a unit root. However, this seems to be due to the low power of the test as the KPSS test cannot generally reject the null hypothesis of stationarity either. If the sample is restricted to exclude the period before February 2002, i.e. before the physical introduction of the Euro, all tests unanimously reject the unit root, albeit some only at 10% significance.

Surprisingly, seasonal adjustment leads to rejecting the null even at 1% in some cases. This is surprising since according to the Maddala and Kim (1998) "the ADF and Philliups-Perron statistics for testing a unit root will be biased towards non-rejection of the unit root null if filtered [i.e. seasonally adjusted] data are used." Nevertheless, the results hold up even after using different information criteria as a robustness check. This might indicate a seasonal unit root in the series, as unit root

[12]Newey-West correction was used together with Bartlett kernel, and Andrews correction with quadratic spectral kernel.

tests on the year-on-year rate reject the unit root more often. However, this may also happen because seasonal effects may contain a bit of the stochastic trend. Adjusting the inflation may get rid of some factors that influence the inflation rate in specific month leading to different results.

The unit root tests for the Euro-area are mixed as well. However, the tests generally reject a unit root in the seasonally adjusted series. Moreover, for inflation in the Euro-area, few KPSS tests reject stationarity. Nevertheless, this result is not very robust as the rejection of stationarity hinges on the inclusion of a trend which is only marginally significant at 10%, and only in small number of cases.

Moreover, based on the economic theory we should also expect inflation to be an integrated process of order zero [I(0)]. An I(0) process is defined as a process which has both finite mean and variance as the number of observations goes to infinity (Verbeek, 2008).

Regarding the mean of the inflation rate, there is only a little doubt that it should be constant, as long as the main objective of the central bank is keeping the price stability. ECB fits this description perfectly as its only mandated objective is price stability[13].

When it comes to the variance of the inflation rate theoretical arguments become bit trickier. Here it is still reasonable to assume that the central bank preoccupied with price stability also cares about keeping the variance of inflation constant. However, this might not be always possible in the face of unexpected and asymmetric macroeconomic shocks. Nevertheless, even allowing for this the series should still be stationary with structural breaks, as given the central banks' objective variance should be at least mean reverting. There is in fact some support for this as Garcia and Perron (1996) provide convincing evidence for structural breaks in mean and variance of inflation in the United States.

This problem becomes even more complex when we consider the possibility of hitting the zero lower bound. There is an ongoing discussion about whether monetary policy can still be effective at the zero lower bound (Krugman et al., 1998). If the monetary policy would become completely impotent, then the ECB would lack the power to control it, making the aforementioned arguments invalid. Nevertheless, this does not seem to be the case. While the conventional policy seems to really loose its bite at the zero lower bound, empirical research shows that central banks can still control inflation through more unconventional monetary policy (Gambacorta et al., 2014, Wu and Xia, 2016).

Furthermore, while the jury might still not be completely out, the assertion that

[13] This objective was further clarified by the Governing Council of ECB as keeping the inflation rate "below, but close to, 2% over the medium term.

inflation is generally an I(0) process is also supported by the more general evidence. While in the past inflation was considered generally to be an I(1) process, mainly due to the highly influential work of Johansen (1992), this notion was later challenged by Culver et al. (1997), Rose (1988) and Basher and Westerlund (2008) who argue that inflation is generally I(0).

3.1.2 Structural Parameters

After the inflation rates, the structural parameters θ_{NL} and θ_{EA} are the most important covariates, and thus they will receive careful treatment here. The most direct way to start getting these structural parameters is without a doubt equation 27 which omitting q_{iT} (as it is equal to 0 by assumption[14]) reads as:

$$\theta_i = (1 - \gamma_i)[a_{iT} - a_{iN} + (b_{iN} - b_{iT})l_i] \tag{32}$$

The parameter γ_i is the fraction of real income spent on tradable goods. However, the subsequent estimations will omit this parameter because data on this parameter are not available[15] .

The a_{ij} denotes the real productivity of industry. There are various ways of measuring the productivity of a particular industry. The most common one is the labour productivity, based on gross value added (GVA). The measure is imperfect, but the data required to calculate the labour productivity are readily available. This is true even at a quarterly frequency which is challenging for the other measures such as KLEMS multifactor productivity (Organisation for Economic Co-operation and Development, 2001).

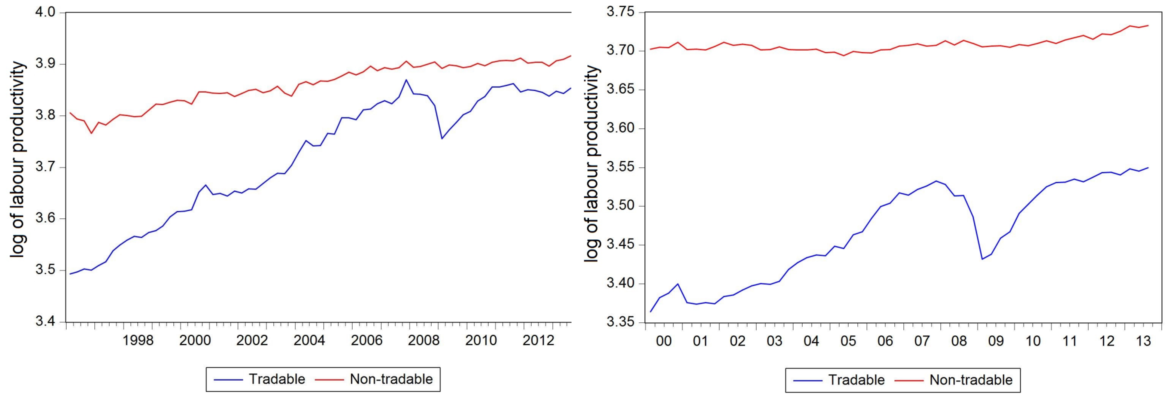
In this work, we will define labour productivity (LP) as the real Gross Value Added divided by hours worked (HW) or formally $LP = \frac{rGVA}{HW}$. The sectoral level data used to calculate this measure for both tradable and non-tradable sectors in the Netherlands and for the Euro-area (excluding the Netherlands) were calculated based on the data from Eurostat[16]. Natural logs of the labour productivity in the Netherlands and the rest of the Euro-area are plotted on the figure 2a and 2b below.

An interesting finding is that productivity in the non-tradable sectors is higher than in the tradable sectors. This holds both in the Netherlands and the rest of the

[14]This is so because under the law of one price log of nominal exchange s equals the difference of logs of price levels p_{NL}, p_{EA}, that is $s = p_{NL} - p_{EA}$. The real exchange rate is by definition $q \equiv s - p_{NL} + p_{EA}$ and thus under the law of one price $q \equiv 0$.

[15]To be more precise, some limited data on the fraction of consumer spending on imports and home production is available. Unfortunately, there is no dataset which provides this data on monthly or at least quarterly basis. Therefore, these datasets are of no use for this research.

[16]Namely from databases with codes: namq_10_a10, namq_nace10_e, and namq_nace10_p.



(a) The Netherlands. **(b)** The Euro-area (Excluding the NL).

Figure 2: Natural log of labour productivity in the Netherlands and Euro-area (Excluding NL). Source: Eurostat's amq_10_a10, namq_nace10_e and namq_nace10_p databases and author's calculations.

Euro-area. However, this is most likely caused by the definitions of tradable and non-tradable sectors, which are described below in appendix C.

This finding can be partially explained by the fact that productivity in the real estate sector might be grossly overestimated. The reason for this is that gross value added is not adjusted for depreciation, which is the major cost that businesses in the real estate sector incur. In practice, this often makes the real estate sector the most productive, in some cases being even twice as productive as the second most productive sector.

For this reason, researchers interested in measuring productivity, such as Elbourne and Grabska (2016) and others, tend to exclude the real estate sector from their measurements. However, this approach might not be best for the purpose of this paper. There are two main reasons for this.

First, the correlation between non-tradable sectors including and excluding the real estate sector is 0.9974 in the Netherlands, and is 0.8595 in the rest of the Euro-area. Hence, removing the real estate sector will only result in a decrease of the mean.

Second, as the figure 26 shows, for the majority of the sample period, the non-tradable sectors still have higher productivity than the tradable ones. This supports the suspicion that this unexpected result is driven by the way how individual sectors are assigned to the tradable and non-tradable sectors. However, at the same time figure 26 suggests that completely removing real estate sector might not be the right solution. Even if we would assume that the correct productivity in the real estate sector is only a third of what we measure, the difference between the tradable and non-tradable would always be negative. Nevertheless, if we exclude it completely there are parts of a sample where this difference becomes marginally positive, and this could affect the signs of the estimated coefficient.

This observation is relevant for the expected sign of some coefficients. For example, the eq. 28 tells us that the price index in the Netherlands should depend positively on the differential of productivity in the Netherlands. Nevertheless, this holds only when the differential is itself positive.

The term $1 - b_{ij}$ represents the output elasticity of labour, which is perhaps the hardest parameter to come by. The reason for this is that detailed micro data which would help to determine these for the Netherlands and the Euro-area (excluding the Netherlands) are either missing or are not comparable. Data from Eurostat's databases[17] are not sufficiently detailed. They only include real gross value added and hours worked for 12 sector which is not enough for proper cross-section estimation.

Estimating these from time series data would be questionable as in the Cobb-Douglas function the elasticities are usually thought of being determined by the available technology (Cobb and Douglas, 1928). Estimating this coefficient from time series data would give us the average expected elasticity across the sample period. Thus using the time series data would require an assumption of constant productivity and technology in each industry during the sample period, which ranges from 1995Q1 to 2013Q3 for the Netherlands, and from 2000Q1 to 2013Q3 for the Euro-area. Hence, using just average across the sampling period would have to exclude all changes and variation in growth of technology during the sample period.

Moreover, in this paper, we are not interested in knowing the elasticity per se, but rather we are interested in knowing it for the sake of estimating the relationship between inflation in the Netherlands and the rest of the Euro-area. Having time invariant estimates of $1 - b_{ij}$ would not add any more variation to the joint term $(b_{ijN} - b_{iT})l_i$, and it would only rescale this term.

The log of labour supply l_i was calculated directly using the data of hours worked. Using hours worked has an advantage of capturing small variations in the labour supply which might not show up when measuring the labour supply using other measures such as employment.

The distinction between the tradable and non-tradable sectors was done based on convention set up by the European Commission's annual macro-economic database (AMECO). The common practice is to regard sectors A, B-E, C and G-I as tradable, and sectors F, J, K, L, M-N, O-Q and R-U as non-tradable (see the table 40 in appendix C for an explanation of the codes.). However, splitting sectors based on

[17]Specifically, databases with codes: namq_10_a10, namq_nace10_e and namq_nace10_p.

the Statistical Classification of Economic Activities in the European Community, commonly known as NACE code, is not necessarily the best practice. Zeugner (2013) argues that the share of value added embodied in foreign demand by sector is a much better way how to measure the tradable and non-tradable sectors, as this allows for splitting all sectors into their respective tradable and non-tradable parts. Nevertheless, this is problematic since data for this share are almost non-existent. The most extensive TiVA database constructed by the OECD provides the share of value added embodied in foreign demand, for the Netherlands and the Euro-area countries, only for 7 years. The figure below plots the share in percent against the time, showing the poor state of the data.

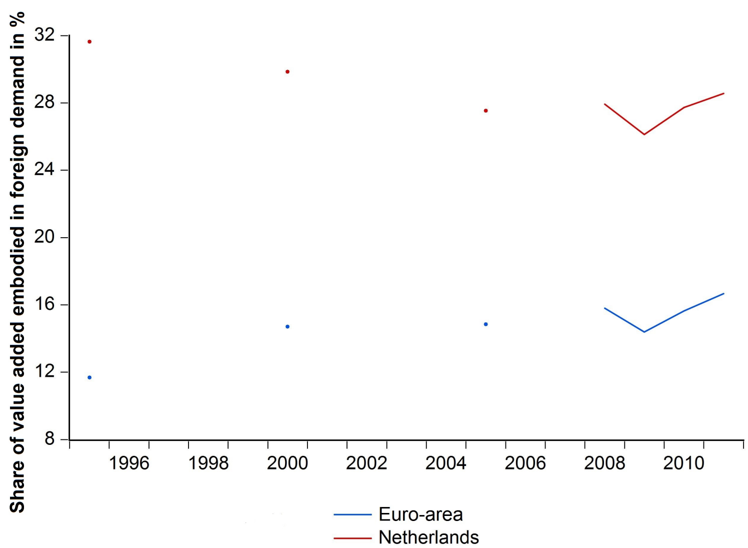


Figure 3: Share of value added embodied in foreign demand in the Netherlands and the EA (measured in %). Source: OECD (2015) TiVA database.

Some of the missing years could be calculated from the World Input-Output Database (WIOD) (Timmer et al., 2015), which contains data for years 1995-2011. However, we would still be missing similar breakdown for the labour supply in each sector, and thus doing this would be to no avail.

Another problem is that almost all of the above-mentioned controls are available only on quarterly frequency. However, this should not be very problematic because all variables used for the calculation of the thetas, such as productivity, factor input elasticities and labour supply, do not vary too much from quarter to quarter, especially after seasonal and calendar adjustment and thus interpolation should not be that far from having actual data on monthly frequency. The interpolation was done using the Catmull-Rom spline, which has several advantages compared to other interpolation methods. First, the Catmull-Rom spline does not overwrite the actual observed values. Second, it avoids the Runge's phenomenon[18] (as opposed to ordinary polynomial interpolation), as it subdivides the sample into small functions which are approximated using a low-degree polynomial. The results of each individual interpolation plotted against the quarterly values can be seen in the appendices D, E and F.

Furthermore, all control variables are tested for the presence of a unit root. These tests were performed using the ADF test with the AIC as a selection criterion for the optimum number of lags[19]. The results of these unit root tests are reported in the table 38 in appendix C.

3.1.3 Control Dummies

The most important control dummies are the tax hikes, as these were found by Mellens et al. (2014) to have a significant impact on prices. Both VAT hike dummies are set to 1 for twelve consecutive months after their implementation and 0 otherwise, for the estimation using the year-on-year rate. In the estimation of monthly annualized rate these dummies will be set to 1 only in the month when the hike occurred.

The rationale for this difference is that year-on-year rate is defined approximately as $ln(x_t) - ln(x_{t-12})$. Thus a level shift in prices will have an impact on the year-on-year rate for the whole year. However, since the annualized monthly rate is defined as $[ln(x_t) - ln(x_{t-1})]12$, the tax hike will show up only as one month increase in the inflation rate. This is very well illustrated by figure 1 which shows year-on-year inflation rates and the annualized rates.

Furthermore, we will add a dummy for the physical introduction of the Euro. The reason for this is that according to the De Nederlandsche Bank (2002) this had a real impact on the increase in prices across the board.

[18] Which is a problem of oscillation at the edges of an interval that occurs when using polynomial interpolation with polynomials of high degree over a set of equispaced interpolation points (Cheney and Light, 2009).

[19] Individual series were tested also using KPSS test and the results always reject unit root at the same or higher confidence level as ADF. The individual tests are not reported for the sake of brevity.

4 Results

This section will present the empirical estimations of model described by the equation 31, shown below for a reminder.

$$\pi_{NL,t} = \alpha + \sum_i \beta_i \pi_{NL,t-1-i} + \sum_i \Lambda_i \pi_{EA,t-i} + \upsilon \vartheta_{NL,t} - \nu \vartheta_{EA,t} - \mu \epsilon_{t-1} + \delta_i z_t + \varepsilon_t \quad (31)$$

This equation can be estimated by a number of different methods, and in this section we will try to estimate it using a range of methods to show that the estimations are robust.

First, we will try to estimate this model by 'short-term' $ARDL(p,q)$ model. We can do this easily by just omitting the error correction term $\mu\epsilon_{t-1}$. However, here we also need to add one caveat. As was shown by Stock and Watson (1988), excluding an error correction term from the model specification, in the presence of cointegration, would lead to a small omitted variable bias. In this case the theoretical model indeed indicates that there should be a long run cointegrating relationship between these variables. However, it might still be worth while to estimate the 'short-run' model. The reason for this is that at the present we have only 15 or 17 years of data available. This might not be enough for the long-run equilibrium to manifest itself in the data.

It is true that this research uses monthly data to increase the number of observations for the estimation purposes. However, while this kind of 'zooming in' on the time series can help with the practical estimation, it might not help that much with detecting cointegration. Moreover, this short run specification will be estimated as an $ARDL(p,q)$ model. Therefore, we will already allow for some dynamics in the model and this should help with avoiding the above-mentioned problems.

Next we estimate error correction model using several different methods for robustness check. First, we will estimate the Engle and Granger (1987) error correction model (subsection 4.4.1). The main reason why we also use the Engle-Granger model is that it is a standard approach to cointegration used in literature. However, the downside of the Engle-Granger approach is that it does not allow for lags in dependent variable in the cointegrating vector which is quite restrictive. The Engle-Granger approach is also sensitive to normalization in dependent variable. Moreover, the Engle-Granger Augmented Dickey-Fuller [EG-ADF] residual test tends to lack power because it ignores information about the dynamic interaction between the variables. Furthermore, the Engle-Granger cointegration model allows for testing of only one cointegrating relationship (Verbeek, 2008).

Second, we will estimate the Johansen (1988, 1991, 2002) error correction model (subsection 4.4.2). There are several advantages to the Johansen error correction model. To start with Johansen model is far less restrictive as it allows for testing multiple cointegration relationships and does not require strict exogeneity assumption. Moreover, it also includes the dynamic interactions between the cointegrated variables (Verbeek, 2008).

Third, we will estimate (Pesaran et al., 2001) unrestricted error correction model (subsection 4.4.3). The advantage of this error correction model is that the whole estimation is done using only one equation. This approach places no restrictions on the cointegrated coefficients.

The last subsection will provide concise summary of these estimations. Therefore, readers who are not interested in all technical details may skip directly to the section 4.5.

4.1 'Short-run' Estimates

The starting place for the short-run estimations will be equation 31 which describes the short-run relationship between the inflation in the Netherlands and the rest of the Euro-area. However, as it was already explained in the previous section, estimating the equation 31 directly is difficult as data required to do so are lacking. In order to alleviate that problem we will approximate the equation 31 as follows:

$$\begin{aligned}\pi_{NL,t} = \alpha + \sum_i \beta_i \pi_{NL,t-1-i} + \sum_i \Lambda_i \pi_{EA,t-i} + \eta \Delta da_{NL,t} + \iota \Delta da_{EA,t} + \\ + \omega \Delta l_{NL,t} + \upsilon \Delta l_{NL,t} + \sum_{i=1}^{3} \delta_i z_i + \varepsilon_t\end{aligned} \tag{33}$$

The first part of the equation, covering the inflation in the Netherlands and the rest of the Euro-area (i.e. $\pi_{NL,t} = \alpha + \sum_i \beta_i \pi_{NL,t-1-i} + \sum_i \Lambda_i \pi_{EA,t-i}$), carries directly from the equation 31. Thus this part does not require separate explanation. As a reminder, the lags of inflation in the Netherlands and the rest of the Euro-area are added to capture the dynamics of the relationship between the two.

The second part consists of the structural parameters described by the equation 27. Specifically, $da_i = a_{iT} - a_{iN}$, which is the difference between the productivity in tradable and non-tradable sectors in both the Netherlands and the rest of the Euro-area. The l_i represents the labour supply measured by the hours worked. Both of them are included as year-on-year or annualized changes (depending on the way dependent variable is measured) as their levels have unit root and are $I(1)$.

The z_i are the three control dummies; 2001 VAT hike, 2002 physical introduction of euro and 2012 VAT hike which were described in the previous section. The δ_i are their respective coefficients.

This model will be estimated in two versions. First version will measure inflation as the year-on-year rate, as well as other control variables for the sake of consistency. The second version will measure inflation and control variables using seasonally adjusted monthly annualized rate.

The equation 33 is essentially an autoregressive distributed lag model (ARDL(p,q)). Insofar, the models can also be estimated putting restrictions on the constant, and the optimal number of lags for inflation can be selected using the information criteria. Most suitable information criterion for this job seems to be the Schwartz information criterion (McQuarrie and Tsai, 1998).

The reason for this is that AIC tends to select too many lags, even if they are insignificant and do not contribute much to the regression (McQuarrie and Tsai, 1998). Indeed in this case AIC selects ARDL(11,7) as the best year-on-year model, and ARDL(6,4) as the best monthly annualized model. This holds despite many of the lags being insignificant. In contrast, using SIC information criterion selects ARDL(1,1) in the year-on-year case and ARDL(1,0) in the estimation which uses annualized rates. These estimations do not include any insignificant lags, but they still have good fit. Moreover, by eliminating lags manually using general-to-specific approach, by eliminating lagged variables until only significant ones remain, we arrive to almost exactly same conclusion as SIC, as this procedure leads to ARDL(1,1) for both year-on-year and annualized rate estimation. Since there are almost no practical differences between ARDL(1,0) and ARDL(1,1) in the annualized rate estimation we will prefer the second one to make year-on-year and annualized rate estimations more comparable.

The output of year-on-year estimation, as well as it's robustness check, can be seen in the table 1. Table 2 reports the estimations using the annualized monthly inflation rate. The first regression model shows the estimation of equation 33[20]. Models 2-6 serve as a robustness checks. The sample covers period from January 2001 to September 2013, as this was the longest sample possible over which data for all covariates were available[21].

The estimations shown in tables 1 and 2 are generally consistent with the economic theory and model presented in the second section, although not unanimously.

[20] Albeit, only one lag of inflation in the Netherlands and the Euro-area was included because further lags were not significant.

[21] However, to provide further robustness checks, table 47 and table 48 include regressions with extended periods in cases where this is possible. For example, due to exclusion of some covariates.

Starting with the inflation in the Euro-area, both signs of year-on-year and monthly annualized rate estimations are consistent with the theoretical predictions and the mean estimates are very close to each other.

The year-on-year mean estimate in the first model is approximately 0.45, with margin of error at 95% confidence interval being between 0.28 and 0.63. The mean estimate from the first model of annualized rate of inflation is 0.52, with margin of error at 95% confidence level laying between 0.37 and 0.7. This indicates that on the average marginal increase in the year-on-year inflation in the rest of the Euro-area will increase the inflation in the Netherlands by about half of a percent. Moreover, this relationship is more or less robust across the various estimations. The only exceptions are the fifth model and the third year-on-year model. In the first case, this is not surprising given that the fifth models does not control for large VAT hikes in the Netherlands. The lower estimate found in the third year-on-year model is harder to waive off, but it can be explained by the fact that it excludes the labour supply in the rest of the Euro-area, which was found to be significant in the first and second model.

However, there are few important caveats pertaining to these estimates. To begin with, these estimates should be interpreted with caution. In the long-run steady state equilibrium and controlling for all structural factors the coefficient should be close to one. Nevertheless, this does not necessarily mean that the found coefficient of roughly 0.5 is inconsistent with theory. The relationship described by the equation 28 implies that price indices (and thus by extension their growth rates as well) should be approximately equal, but this holds only for long run equilibrium[22] and is conditional on all other structural factors. Also, the period covered in this study, which ranges from January 2001 to September 2013 was a very turbulent one. Moreover, there are most certainly many structural factors that are hard to control for. There are many market imperfections which also disturb this relationship. Because of these the relationship may be permanently different from unity.

For example, if the levels and changes of direct or indirect taxes are different between the Netherlands and the rest of the Euro-area, so will the price levels and inflation rates. Moreover, this caveat is not restricted to taxes only as ultimately almost any difference in government policies might have similar effect, as many government regulations may implicitly affect price levels and their respective growth rates. Differences in policies ranging from minimum wage via social programmes to plethora of regulations may cause similar deviations from the relationship in a very similar manner to taxation. Unfortunately, it is not possible to control for all these

[22]Indeed, the estimations which utilize longer time periods, shown in table 47 and 48, have on average higher coefficient for inflation in the Euro-area.

differences. It is reasonable to assume that as policies between the Netherlands and the rest of the Euro-area converge, so will the relationship and vice versa[23].

Furthermore, the estimations presented in the tables 1 and 2 are not direct representations of the equation 30. This is so because the changes in the productivity should also be weighted by the changes in income spent on tradable goods. This should also be applied to the changes in labour supply which should be additionally weighted by the differences of factor-input elasticities in non-tradable and tradable sectors. However, as it was mentioned previously these are incredibly hard to measure consistently and on sufficient frequency. Thus they could not be included in the estimations.

Insofar, it is possible that many of these unaccounted differences 'feed into' the estimates. This does not mean that these coefficients are useless, as these might be still of great interest to the practical policy makers. However, estimations of these coefficients should be interpreted with these caveats in mind.

The first lags of inflation in the Netherlands and the rest of the Euro-area serve to capture dynamics in the series. Note that the tables report these variables using the lag operator $L^n x = x_{t-n}$ to conserve space.

When it comes to the year-on-year and monthly annualized differentials in productivity between tradable and non-tradable sectors in the Netherlands and the rest of the Euro-area, the estimations from the tables 1 and 2 are mostly insignificant. Moreover, they also have opposite signs from what we would expect based on the model presented in the second chapter in the year-on-year estimations.

However, there are several factors that might explain this result. First, as the table 3 shows most of the variables suffer from very high multicollinearity. Multicollinearity might lead to poorly estimated coefficients. Especially when the estimated coefficients are close to being zero, such as in this case, the multicolinearity might even cause signs of coefficients to flip. This might also explain why the signs are so different between the year-on-year estimations and the annualized rate estimations. The table 3 shows that variance inflation factors are lower in the estimation that uses annualized rates.

Furthermore, as it was mentioned in the previous sections, measuring productivity is very hard, and in addition also assigning sectors as tradable and non-tradable solely on their NACE code is also not very proper as well.

Another two important controls are labour supply in the Netherlands and the rest of the Euro-area. When it comes to the labour supply it is very hard to tell what sign it should have as the equation 28 says that this will depend on the relative

[23]Although such convergence is not an explicit goal of the European Union, it will be necessary amongst the Euro-area members in order to manage macroeconomic shocks better.

factor input elasticities in the tradable and non-tradable sectors.

Thus, the relationship can be positive if the difference between the elasticity in non-tradable sectors is higher than in the tradable, or negative in the opposite situation. Furthermore, in the case that the factor input elasticities are equal in the tradable and non-tradable sectors the coefficient should be zero. Insofar, here even an insignificant result might imply that the factor input elasticities are equal between the tradable and non-tradable sectors in a particular country.

In table 1, we can see that the change in the hours worked in the Netherlands has positive impact on the change in inflation, but it is insignificant. Nevertheless, this might again be result of multicollinearity as removing the productivity differentials increases its t-statistics considerably,from about 0.93 to about 1.57, almost making it significant at the 10% confidence interval.

The estimated coefficient for the change in labour supply in the rest of the Euro-area is negative and it is significant in the first year-on-year model, although only at 10% significance interval.However, after removing the productivity differentials it becomes highly significant even at 1% confidence interval.

The coefficient has value of approximately 0.10. This means that on an average, marginal positive change in the log of labour supply in the rest of the Euro-area decreases Dutch inflation by about one tenth of a percent. Nevertheless, one should not read into the value of this coefficient by much. The equation 28 shows that log of the labour supply and the difference between factor input elasticities enter the equation multiplicatively. However, since we are unable to observe the factor input elasticities directly, it is probable that part of the coefficient reflects also the difference between those two. Thus the coefficient likely reflects both changes in these elasticities and the log of labour supply together, with exception of special case where the difference between those two would always be equal to one, which is extremely unlikely.

Table 2 shows that the log of labour supply is always insignificant in the annualized rate estimations. However, this might be also caused because the change in the labour supply is more pronounced when measured as a year-on-year change, as it is not unreasonable to assume that hours worked do not fluctuate much each month, in the Europe which has traditionally stronger labour protection laws and labour market regulation[24].

The 2001 VAT increase from 17.5% to 19% was found to be highly significant across all estimations. In the year-on-year estimation first two models estimate that the 2001 tax hike increased year-on-year inflation by about 1.5 percent[25]. This, is

[24]Remember that the series are already adjusted both seasonally and for the calendar.

[25]Note that since the rates used in both estimations were in decimals, coefficient of level variables

consistent with the research of Mellens et al. (2014) which shows that the 2001 VAT hike was fully passed onto prices.

In the annualized inflation rate estimation, this VAT hike resulted in a sharp and highly significant spike of about 21% in January 2001 [26]. This is also visible on the graph 1. Again, this is generally consistent with the findings of Mellens et al. (2014), as they too find that the pass-through to prices was immediate without any anticipation effects, and large spike in monthly inflation is exactly what one would expect to find in such a case.

Physical introduction of the euro currency seemed to have significantly increased prices in the Netherlands by almost half of a percentage point. This is slightly higher than De Nederlandsche Bank (2002) estimate of roughly one quarter of percentage point. Nevertheless, De Nederlandsche Bank (2002) estimate is still within 95% confidence interval. The annualized rate of inflation increased about 10% at the beginning of 2002[27]. Both of these support the assessment of De Nederlandsche Bank (2002), which reported that the physical introduction of the euro was used by businesses to increase prices.

The 2012 VAT hike from 19 to 21% also significantly increased the inflation rate. However, it increased the year-on-year rate only by less than half of a percent, and the annualized rate only by about 10%. This is lesser than expected. However, this can be explained by the fact that as the figure 1 shows, whole currency area was sliding into deflation since the 2012, following the infamous premature interest hike by ECB and the Greek debt crisis.

There are also some notable differences in the goodness of fit and standard errors between the year-on-year and annualize rate model. These are most likely caused by the fact that the annualized rate varies about twice as much as the year-on-year rate. These differences are explored in more detail in the subsequent chapters.

must be pre-multiplied by 100 to get to percentage.

[26]This would represent roughly 1.75% increase across the year. This is a bit higher than 1.5% VAT increase as it also includes ‘normal’ increase in the price level that month.

[27]This would translate to roughly 0.8% yearly increase.

Table 1: Summary of the estimations of the relationship between the year-on-year inflation rate in NL and EA.

Overview of Estimations (Year-on-Year Rate). Sample:01/2001-09/2013.						
Models:	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	π_{NL}	π_{NL}	π_{NL}	π_{NL}	π_{NL}	π_{NL}
π_{EA}	0.452344***	0.442506***	0.311526**	0.517058***	0.339936*	0.379571***
	[0.275, 0.630]	[0.318, 0.567]	[0.076, 0.547]	[0.275, 0.759]	[-0.038, 0.718]	[0.131, 0.628]
	(5.040777)	(7.034410)	(2.611446)	(4.223157)	(1.777482)	(3.016931)
$L^1\pi_{EA}$	-0.206517**	-0.206872**	-0.130525			-0.153492
	[-0.395, -0.018]	[-0.389, -0.025]	[-0.399, 0.138]			[-0.430, 0.123]
	(-2.168909)	(-2.242176)	(10.16986)			(-1.098314)
$L^1\pi_{NL}$	0.710984***	0.718548***	0.740147***			0.705947***
	[0.584, 0.838]	[0.583, 0.854]	[0.596, 0.884]			[0.559, 0.853]
	(11.07338)	(10.49728)	(10.16986)			(9.494059)
Δda_{NL}	-0.005726					-0.024845
	[-0.052, 0.041]					[-0.061, 0.011]
	(-0.244775)					(0.010810)
Δda_{EA}	0.001317					0.007080
	[-0.027, 0.029]					[-0.025, 0.039]
	(0.093597)					(0.439476)
Δl_{NL}	0.033805	0.039533				
	[-0.038, 0.106]	[-0.010, 0.089]				
	(0.929783)	(1.574959)				
Δl_{EA}	-0.098653*	-0.105260***				
	[-0.205, 0.008]	[-0.181, -0.030]				
	(-1.832923)	(-2.754901)				
Dummy 2001 VAT	0.015664***	0.015644***	0.010535***	0.034333***		0.011724***
	[0.008, 0.024]	[0.008, 0.024]	[0.004, 0.017]	[0.030, 0.039]		[0.005, 0.018]
	(3.796509)	(3.910122)	(3.255470)	(5.575281)		(3.546389)
Dummy 2002 Euro	0.004725***	0.004622***	0.004161**	0.020591***		0.004728**
	[0.001, 0.008]	[0.001, 0.008]	[0.001, 0.008]	[0.013, 0.028]		[0.001, 0.008]
	(2.738975)	(2.655585)	(2.337731)	(5.575281)		(2.576880)
Dummy 2012 VAT	0.003360***	0.003206***	0.004367***	0.015493***		0.004733 ***
	[0.001, 0.005]	[0.002, 0.005]	[0.002, 0.007]	[0.012, 0.020]		[0.002, 0.007]
	(3.206850)	(3.836070)	(3.494928)	(7.598525)		(3.787104)
Constant	0.000217	0.000256	0.000646	0.006409**	0.015522***	0.000531
	[-0.002, 0.002]	[-0.002, 0.002]	[-0.001, 0.002]	[0.001, 0.020]	[0.006, 0.025]	[-0.001, 0.002]
	(0.226221)	(0.293252)	(0.734532)	(2.262770)	(3.153482)	(0.559018)
Adj. R^2	0.954450	0.955032	0.951401	0.833031	0.043100	0.951936
F-Stat.	319.4992	404.5235	496.9349	190.5865	7.846317	377.3025
SE of reg.	0.002554	0.002538	0.002638	0.004890	0.011707	0.002624
$\sum \varepsilon^2$	0.000926	0.000927	0.001016	0.003539	0.020693	0.000991
Durbin-Wat. stat.	1.520510	1.533698	1.511693	0.363121	0.050338	1.465556
Wald F-stat.	589.4569	661.5608	668.9513	592.7518	3.159443	472.3445
Errors:	HAC	HAC	HAC	HAC	HAC	HAC

The μ and σ of the dep. var. druing the sample period were 0.022414 and 0.011967 respectively.
HAC standard errors & covariances were calculated using prewhitened standard errors,
Bartlett kernel and Newey-West bandwidth.
t-statistics in (). Lower (L) and Upper (U) 95% confidence intervals are included in [L, U].
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 2: Summary of the estimations of the relationship between the annualized inflation rate in NL and EA.

Overview of Estimations (Anualized Rate). Sample: 01/2001-09/2013.						
Models: Dependent Variable	(1) π_{NL}	(2) π_{NL}	(3) π_{NL}	(4) π_{NL}	(5) π_{NL}	(6) π_{NL}
π_{EA}	0.532073***	0.491295***	0.488013***	0.466842***	0.157412	0.528494***
	[0.369, 0.696]	[0.344, 0.639]	[0.341, 0.635]	[0.303, 0.631]	[-0.423, 0.737]	[0.357, 0.700]
	(6.428748)	(6.578845)	(6.567948)	(5.620174)	(0.536229)	(6.094365)
$L^1\pi_{EA}$	-0.084534*	-0.127623***	-0.132350***			-0.089831**
	[0.064, 0.126]	[-0.218, -0.037]	[-0.218, -0.047]			[-0.177, -0.003]
	(-1.691215)	(-2.794740)	(-3.064609)			(-2.048534)
$L^1\pi_{NL}$	0.063832**	0.093806***	0.096307***			0.066747**
	[0.002, 0.126]	[0.032, 0.156]	[0.039, 0.153]			[0.011, 0.123]
	(2.031415)	(3.001487)	(3.336713)			(2.356294)
Δda_{NL}	0.021733					0.023233
	[-0.030, 0.074]					[-0.027, 0.074]
	(0.824308)					(0.912640)
Δda_{EA}	-0.107594***					-0.108767***
	[-0.177, -0.038]					[-0.173, -0.045]
	(-3.070216)					(-3.374065)
Δl_{NL}	-0.041739	-0.041980				
	[-0.107, 0.024]	[-0.121, 0.037]				
	(-1.264407)	(-1.052760)				
Δl_{EA}	0.008856	0.007646				
	[-0.054, 0.071]	[-0.060, 0.075]				
	(0.280352)	(0.224600)				
Dummy M	0.211637***	0.213099***	0.212626***	0.209776***		0.211371***
2001 VAT	[0.197, 0.226]	[0.197, 0.230]	[0.197, 0.228]	[0.193, 0.227]		[0.196, 0.227]
	(28.70419)	(25.30594)	(27.36040)	(24.09229)		(27.17699)
Dummy M	0.097335***	0.097226***	0.098175***	0.100358***		0.098291***
2002 Euro	[0.091, 0.103]	[0.093, 0.102]	[0.094, 0.103]	[0.096, 0.105]		[0.093, 0.104]
	(34.29585)	(43.47245)	(43.65817)	(48.01236)		(36.07639)
Dummy M	0.102481***	0.106762***	0.106399***	0.104700***		0.102068***
2012 VAT	[0.099, 0.106]	[0.103, 0.111]	[0.103, 0.110]	[0.101, 0.108]		[0.099, 0.105403]
	(55.05178)	(49.21346)	(56.73005)	(61.48706)		(60.49526)
Constant	0.009639***	0.009728***	0.009792***	0.009716***	0.018514**	0.009716***
	[0.004, 0.015]	[0.005, 0.014]	[0.005, 0.014]	[0.101, 0.108]	[0.003, 0.034]	[0.004, 0.015]
	(3.610311)	(4.234178)	(4.422463)	(3.695859)	(2.415954)	(3.618817)
Adj. R^2	0.500529	0.487750	0.493461	0.484173	0.005820	0.506182
F-Stat.	16.23222	19.09128	25.67928	36.66807	1.889843	20.47569
SE of reg.	0.019848	0.020100	0.019988	0.020170	0.028002	0.019735
$\sum \varepsilon^2$	0.055938	0.058177	0.058328	0.060211	0.118400	0.056084
Durbin-Wat. stat.	1.936766	1.900934	1.902522	1.728015	1.479855	1.937852
Wald F-stat.	N/A	N/A	N/A	N/A	0.287542	N/A
Errors:	HAC	HAC	HAC	HAC	HAC	HAC

The μ and σ of the dep. var. druing the sample period were 0.021610 and 0.028084 respectively.
HAC standard errors & covariances were calculated using prewhitened standard errors,
Bartlett kernel and Newey-West bandwidth.
t-statistics in (). Lower (L) and Upper (U) 95% confidence intervals are included in [L, U].
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 3: Variance Inflation Factors (VIF) of covariates used for year-on-year and annualized monthly rate estimations.

VIF from year-on-year estimations.							
Variables:	π_{EA}	$L^1\pi_{EA}$	$L^1\pi_{NL}$	Δda_{NL}	Δda_{EA}	Δl_{NL}	Δl_{EA}
coeff. variance	0.008053	0.009066	0.004122	0.000547	0.000198	0.001322	0.002897
Centered VIF	9.230909	10.26302	25.41605	14.53722	6.599831	5.646191	39.04765

VIF from monthly annualized estimations.							
Variables:	π_{EA}	$L^1\pi_{EA}$	$L^1\pi_{NL}$	Δda_{NL}	Δda_{EA}	Δl_{NL}	Δl_{EA}
coeff. variance	0.082036	0.016159	0.006610	0.002820	0.004720	0.003222	0.019072
Centered VIF	5.733694	3.142136	4.682377	4.940566	4.074230	1.219652	5.537682

4.2 Notes on the Fit and Residual Diagnostics of Models

Some readers might be worried about high R^2 of year-on-year models, which exceeds 0.95 in some of them. While usually higher R^2 is preferred, high R^2 is also one of the symptoms of endogeneity (Verbeek, 2008). However, in this case the high R^2 is achieved mainly by accounting for dummies. This can be seen from comparing the 5th model to the other models, which shows that the R^2 drops considerably after removing dummies from the regression. Indeed the models 4 and 5 were estimated mainly to show this.

The heteroskedasticity in the main models was tested using the White test. White test (White, 1980) was chosen because it is more broad and it does not require explicit assumption about the form of heteroskedasticity, even though this comes at the price of a bit lower power.

The table 4 shows that the test finds heteroskedasticity in the year-on-year estimation according to all the test statistics. However, in case of the annualized monthly rate, test is ambiguous as F-statistics and NR^2 cannot reject the null of homoskedasticity. On the other hand, the normalized sum of squared residuals suggests that there is some limited evidence of heteroskedasticity. Additionally, plot of residuals shown on figure 29 shows that there really might be some heteroskedasticity present especially later in the series. These results are generally similar across the models.

After discussing briefly the heteroskedasticity tests, we can turn our attention to testing for autocorrelation. Usually the first test statistics that people turn to for autocorrelation is the Durbin-Watson statistics. With 153 observations and 10 regressors the lower and upper limit for the indeterminate range, according to Savin and White (1977), is 1.51 and 1.75 respectively. According to this criterion most of the models do not suffer from first order auttocorrelation, as the values lie either above or within the indeterminate region. However, Durbin-Watson statistics is un-

Table 4: Main results from the White test on residuals from the first model of year-on-year and annualized rate estimations.

White Test on Residuals from the First Year-on-Year Model

F-statistic	2.406418	Prob. F(29,123)	0.0005
Obs*R-squared	55.38406	Prob. Chi-Square(29)	0.0022
Scaled explained SS	139.2254	Prob. Chi-Square(29)	0.0000

White Test on Residuals from the First Monthly Annualized Rate Model

F-statistic	0.828254	Prob. F(38,114)	0.7431
Obs*R-squared	33.10201	Prob. Chi-Square(38)	0.6952
Scaled explained SS	310.9637	Prob. Chi-Square(38)	0.0000

reliable in models with lagged dependent variable. Removing the lagged inflation from the regression decreases the Durbin-Watson statistics substantially, way below the indeterminate region. Thus, we can be sure that there is first order autocorrelation in the models without the lagged dependent variable, but just based on the Durbin-Watson it impossible to judge whether explicitly adding lag of inflation in the Netherlands solved this issue or not. Insofar, the series will be additionally tested with the Breusch (1978)-Godfrey (1978) Serial Correlation LM test.

The Berush-Godfrey LM test does not find autocorrelation in the first year-on-year and monthly annualized rate models[28]. This is a good sign as Verbeek (2008) argues that in the time series estimations, tests for autocorrelation can be interpreted as misspecification tests. Thus lack of finding autocorrelation in the first model, but finding it in the subsequent models, without the lags of dependent variable, means that these models are as parsimonious as we can get without being misspecified.

Table 5: Main results from the serial correlation LM tests on residuals from the first year-on-year and annualized monthly rate models.

Breusch-Godfrey Serial Correlation LM Test for the first year-on-year model:

F-statistic	1.469087	Prob. F(12,130)	0.1438
Obs*R-squared	18.27041	Prob. Chi-Square(12)	0.1077

Breusch-Godfrey Serial Correlation LM Test for the first monthly annualized rate model:

F-statistic	1.695680	Prob. F(12,130)	0.0747
Obs*R-squared	20.70706	Prob. Chi-Square(12)	0.0548

To deal with the problems of heteroskedasticity we will use Whitney K. Newey (1987) (HAC) standard errors. Following the Newey and West, errors were first

[28]Moreover, this is generally true for models that also include the short term dynamics.

prewhitened where a number of lags were based on Schwartz criterion. The maximum number of lags were determined based on the sample size. Second, the data dependent bandwidth parameter was estimated using Bartlett kernel and Newey-West automatic bandwidth method.

4.3 Addressing Further Endogeneity concerns

Some might worry about endogeneity issue since European Central Bank [ECB] targets the Euro-area-wide inflation including the Dutch one. However, this will not be an issue as long as we can think of the Netherlands as (economically) small country vis-à-vis the rest of the Euro-area.

Also the model based on Ca'Zorzi et al. (2005) sketched above presupposes that the Netherlands is a small country, while the Euro-area can be thought of as a big country. Nevertheless, the Dutch economy is assigned only a negligible weight in the basket which determines the country's contribution to the Euro-area inflation (5% on average[29]). While there are countries which have even smaller weight (for example Estonia, Cyprus, Latvia, Lithuania, Luxembourg and Slovenia, that do not get even one full percentage point on average), the 5% weight is still negligible compared with larger Euro-area countries (for example, the average weight of Germany, France, Italy and Spain is 29.11%, 20.49%, 18.17% and 10.93% respectively).

The Netherlands is also smaller compared to the rest of the Euro-area in other economic indicators as well. For example, the hours worked in the Netherlands also represent, on average, only about 5.11% of hours worked and 6.68% of Real Gross Value added from the whole Euro-area.

In fact 5% weight put on inflation seems like a reasonable threshold for assuming that the inflation in particular a country does not affect the ECB monetary policy. Insofar, the inflation in the Netherlands can be thought of as exogenously determined by the inflation in the Euro-area and the differences in the structural factors.

Nevertheless, it must be said that 5% is of course only arbitrary threshold. Moreover, it could also be possible that there might be complex political forces which would compel ECB to put more weight on the Netherlands than justified by the size of its economy, but this seems to be unlikely.

To settle this we can asses the possibility of reverse causality directly from the data by running the Granger (1969) causality test. The Granger causality test looks whether lags of one variable in a VAR are useful in predicting the other beyond the information contained in its own lags. Even though Granger causality does not imply

[29] The weights were retrieved from the Eurostat's database with code: prc_hicp_cow.

causality in the common sense of the term, in case that both variables are Granger causing each other there might be a reason to worry about endogeneity.

We start by determining the optimum number of lags for the auxiliary VAR used for the Granger causality test. The table 6 presents summary of optimum lag for the year-on-year and annualized rate version of the auxiliary VAR. In the year-on-year version most criteria support 12 lags whereas in the annualized rate version most of them justify only two. Nevertheless, as a robustness check, table 7 reports results of the Granger causality test considering all possible lag values.

The Granger causality test consistently cannot reject the null of non-Granger causality of inflation in the Netherlands on inflation in the rest of the Euro-area. Moreover, with the exception of year-on-year rate test using only one lag, as selected by Schwartz information criterion, the tests can reject the null hypothesis that inflation in the rest of the Euro-area does not Granger cause inflation in the Netherlands. All in all these results are consistent with the assertion that the endogeneity should not be a problem between the dependent and main independent variable.

Table 6: Optimal lag selection for Granger causality test between inflation in the Netherlands and rest of the Euro-area.

Year-on-Year

Lag	LogL	LR	FPE	AIC	SC	HQ
0	1392.611	NA	1.05e-08	-12.69964	-12.66869	-12.68714
1	1868.645	939.0266	1.40e-10	-17.01046	-16.91761*	-16.97296
2	1877.109	16.54026	1.35e-10	-17.05122	-16.89647	-16.98872*
3	1883.646	12.65752	1.32e-10	-17.07439	-16.85774	-16.98690
4	1884.405	1.455463	1.36e-10	-17.04480	-16.76624	-16.93230
5	1889.434	9.552481	1.34e-10	-17.05419	-16.71374	-16.91669
6	1890.061	1.179751	1.39e-10	-17.02339	-16.62103	-16.86089
7	1892.625	4.776612	1.41e-10	-17.01027	-16.54602	-16.82277
8	1895.129	4.619664	1.42e-10	-16.99661	-16.47046	-16.78411
9	1897.852	4.972269	1.44e-10	-16.98495	-16.39689	-16.74745
10	1901.064	5.808971	1.45e-10	-16.97775	-16.32780	-16.71525
11	1914.070	23.27955	1.34e-10	-17.06000	-16.34814	-16.77250
12	1920.037	10.57271*	1.32e-10*	-17.07797*	-16.30421	-16.76547

Annualized Monthly Rate

Lag	LogL	LR	FPE	AIC	SC	HQ
0	985.3932	NA	7.40e-07	-8.441143	-8.411521*	-8.429198
1	994.3790	17.74018	7.09e-07	-8.483940	-8.395072	-8.448104
2	1004.236	19.29108	6.74e-07*	-8.534215*	-8.386102	-8.474489*
3	1004.824	1.139944	6.94e-07	-8.504924	-8.297566	-8.421308
4	1010.950	11.77852	6.82e-07	-8.523172	-8.256568	-8.415666
5	1011.341	0.745495	7.03e-07	-8.492196	-8.166346	-8.360799
6	1015.871	8.554780	7.00e-07	-8.496746	-8.111652	-8.341459
7	1018.840	5.555239	7.06e-07	-8.487894	-8.043554	-8.308717
8	1019.425	1.084489	7.27e-07	-8.458580	-7.954995	-8.255512
9	1020.994	2.882242	7.43e-07	-8.437714	-7.874883	-8.210756
10	1022.935	3.533271	7.56e-07	-8.420045	-7.797970	-8.169197
11	1027.431	8.103211	7.53e-07	-8.424297	-7.742976	-8.149558
12	1039.590	21.70891*	7.03e-07	-8.494332	-7.753766	-8.195703

* indicates lag order selected by the criterion

LR: sequential modified LR test statistic (each test at 5% level)

FPE: Final prediction error

AIC: Akaike information criterion

SC: Schwarz information criterion

HQ: Hannan-Quinn information criterion

Table 7: Summary of Pairwise Granger causality Tests. Sample: 01/1996 - 06/2016

	Year-on-Year		
Lags:	(1)	(2)	(12)
Observations:	(230)	(229)	(219)
	Dependent Variable: π_{NL}		
F-stat.	2.37531	2.58178*	2.21138**
	Dependent Variable: π_{EA}		
F-stat.	0.32303	0.56293	1.39410

	Annualized Monthly Rate		
Lags:	(0)	(2)	(12)
Observations:	(N/A)	(243)	(233)
	Dependent Variable: π_{NL}		
F-stat.	N/A	3.69096**	2.34605***
	Dependent Variable: π_{EA}		
F-stat.	N/A	1.02099	0.76434

F-statistics is from the Wald test for joint significance of all lags from the VAR.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

4.4 Cointegration and Error Correction

To further improve on the results from the previous section we will consider adding an error correction term. Moreover, the error correction term can reveal potentially interesting information about the speed of adjustment of the inflation to its long run equilibrium. Equation 28 suggests that there should also be a long run relationship between the Dutch and the Euro-area's inflation, and the Dutch and the Euro-area's structural factors.

Normally, in situations such as these building an error correction term would be easy exercise as the relationship between the Dutch and the Euro-area's price levels, described by the equation 28, can be easily rearranged to solve for the error correction. To see this lets assume that equation 28 holds in every period, and solve for ϵ. Thus in period $t-1$ we get:

$$\epsilon_{t-1} = p_{NL,t-1} - p_{EA,t-1} - \theta_{NL,t-1} + \theta_{EA,t-1} \tag{34}$$

This could be substituted into equation 33 which would give us the following error correction model

$$\begin{aligned} \pi_{NL,t} = \alpha + \sum_i \beta_i \pi_{NL,t-1-i} + \sum_i \Lambda_i \pi_{EA,t-i} - \mu(p_{NL,t-1} - p_{EA,t-1} \\ -\theta_{NL,t-1} + \theta_{EA,t-1}) + \upsilon_i \vartheta_{NL,t} - \nu_i \vartheta_{EA,t} + \delta_i z_t + \varepsilon_t \end{aligned} \tag{35}$$

Unfortunately, as it was already emphasized several times estimating this equation is currently very challenging, because calculating the θ's and ϑ's requires knowing the factor input elasticities as well as the share of income spent on non-tradables, as well as how these change over time.

Even though we don't have enough data to calculate θ's and ϑ's we can still continue using the productivity differentials and labour supply. However we cannot impose coefficients values anymore on the error correction term. These will have to be estimated from the data with cointegrating regression.

The model selection from the previous section allows for multiple approaches to cointegration and error correction. First, there is the classical Engle and Granger (1987) approach to cointegration and error correction. Results of this approach can be found in the section 4.4.1. Second, we will use Johansen (2002) approach to cointegration and error correction. The Johansen method will be presented in the section 4.4.2. Third, our use of ARDL model allows for using unrestricted error correction of Pesaran et al. (2001). This method will be pursued in the section 4.4.3.

All these approaches have their vices and virtues. Each of these will be described in greater detail in each individual section. By utilizing all of these methods we can see whether the results are robust across the model given their pros and cons.

4.4.1 Engle-Granger approach

The Engle and Granger (1987) approach to cointegration and error correction involves running auxiliary cointegrating regression, and subsequently testing the residuals for unit root. In the case that the residuals are stationary we can use the lag of them as an error correction term.

The auxiliary regression will be built upon the equation 28. However, as in previous section we need to tweak the equation 28 a bit because we are not able to observe all necessary parameters for calculating the structural terms (the thetas). Thus in our 'long run' cointegrating equation will use the productivity differentials and labour supply directly without adjusting them both for the share of income spent on non-tradables and without adjusting the labour supply for the factor input elasticities.

After all adjustments our cointegrating equation will take the following form:

$$p_{NL,t} = \alpha + \beta p_{EA,t} + \eta da_{NL,t} + \iota da_{EA,t} + \omega l_{NL,t} - \upsilon l_{EA,t} + \epsilon_t \tag{36}$$

Furthermore, here we will restrict our sample to only include data from January 2002 onwards. The reason for this restriction is the large shifts in levels of the Dutch price levels, caused mainly by the large 2001 VAT hike.

In principle these could be accounted for by adding appropriate dummies for the shift in the levels. However, in such cases the whole process is even more complicated as the standard MacKinnon et al. (1996) critical values cannot be reliably applied. This is especially true when several breaks need to be accounted for. Moreover, the test's results can quite vary depending on the exact nature of the structural breaks (Arranz and Escribano, 2000). This is problematic mainly when there are more than a one structural break, and restricting the sample after the period of physical introduction of euro will reduce number of structural breaks to account for.

Indeed the Engle-Granger test generally finds no cointegrating relationship, in cases when this level shift is not accounted for. In the case when we account for it by adding additional deterministic dummies for these level shifts, tests find cointegration but their lagged values are positive. Significant and positive error correction terms usually imply a structural break (Verbeek, 2008) or a misspecified cointegrated equation. However, as soon as we will restrict our sample to January 2002 onwards, we can find both signs of cointegration and error correction terms which are within proper bounds.

Although cointegrated equation is consistent and converges at the higher rate than is standard (Hamilton, 1994), the standard OLS estimates are generally non-Gaussian and suffer from several other shortcomings. Because cointegrating equations are know to have non-standard asymptotic distribution we cannot use OLS, because the test statistics would be wrong, and it cannot be saved even with the HAC errors.

These problems, caused mainly by the long run correlation between the cointegrating equation and stochastic regressors innovations, can be avoided using an alternative estimator. One such estimator was created by Phillips and Hansen (1990), who use semi-parametric correction to avoid the above mentioned problems. Insofar, we will employ the Philips and Hansen's estimator, which is also known as fully modified OLS (FMOLS).

We will start by estimating the model described by equation 36, and also two other models which serve as robustness checks. The table 8 shows the FMOLS estimates of the cointegrating equation based on equation 36. As we can see in the table 8 the Euro-area's price level has a statistically significant coefficient ranging from 0.89 to 0.77 across the three estimated models. This is not that far from what we would expect based on the theory presented in the section 2, although coefficients that would be even closer to unity would be more appealing. Nevertheless, this is most likely again just reflecting all the subtle differences between the Netherlands and the rest of the Euro-area which affect overall price level. These include the differences in minimum wages or pensions, and way how these are changing each year, or various other policies.

Unfortunately other coefficients are less robust. The productivity differential in the Netherlands has statistically significant negative sign with coefficient of approximately -0.15 in the first model. In the second model it turns statistically insignificant with value of -0.03. The coefficient of productivity differential in the rest of the Euro-area was positive with the estimated coefficient being roughly 0.09, but also statistically insignificant and thus we cannot reject the hypothesis that the true coefficient is zero.

One might be puzzled by the negative sign of the coefficient in the Netherlands and positive in the rest of the Euro-area, as based on equation 27 one would expect opposite result. However, such result is not inconsistent with the theory. The table 26 shows that the difference between productivity in the tradable and non-tradable sectors was negative but the difference was also narrowing. In such a case the signs would also change in equation 27.

Turning our attention to the labour supply we can see that both, hours worked in the Netherlands and hours worked in the rest of the Euro-area, have negative and statistically significant coefficients of -0.14 and -0.16 respectively. The negative coefficient of hours worked in the Netherlands could imply that the factor input elasticity in the Netherlands is higher in the tradable sectors than non-tradable.

The coefficients are again very small and close to zero, and become statistically insignificant in the robustness check. However, it is well possible that their coefficients are truly zero, as in the long run both tradable and non-tradable factor inputs might become equally (infinitely) elastic which, according to equation 27, would mean that they will drop out from the relationship.

As we discussed before, the Mellens et al. (2014) find that the 2012 VAT hike had a significant impact on the price level in the Netherlands. Because of this we will also allow for this level shift in the price level in the cointegrating equation. As we can see in the table 8 according to our estimations the coefficient was always statistically significant with coefficient of around 0.015 in the first model, 0.016 in the second and 0.027 in the last one.

The long run variance will be estimated using prewhitened standard errors, with optimum lag based on Schwartz criterion and maximum considered lag based on the number of observations. Furthermore, Bartlett kernel and Newey-West automatic bandwidth selection was used.

Because the estimates of the cointegrating equation are not robust to excluding insignificant productivity differential in the Euro-area, we will estimate the ECM with lagged residuals from both model 1 and 3 as a robustness check. The methodology used in these estimations is the same as for estimations in the previous section, so it will not be repeated here. Moreover, since the error correction models are estimated using the 01/2002-09/2013 sub-sample, we will also estimate these models without the error correction term to make them more comparable[30].

The table 9 shows the estimations of the year-on-year error correction model, and the table 10 presents the monthly annualized rate estimations. In both cases, the first model is just an estimation of the model from the previous chapter. This serves as both robustness check and to see how other coefficients change when we add the error correction term. Here we can see that the main coefficient of interest, the effect of inflation in the Euro-area on the inflation in the Netherlands is still robust both across all estimations in this sub-sample as well as in the previous estimates.

However, this does not generally hold for other covariates. The signs and values are still very similar, with few exceptions, but most of the covariants are now insignificant.

Focusing on the error correction term, table 9 shows that error correction in the

[30]Although, because model uses lagged variables this practically means that the sample will be restricted to 02/2002-09/2013.

Table 8: Summary of the estimations of cointegrating relationships.

Overview of the FMOLS estimates of the cointegrating relationship. Sample: 01/2002-09/2013			
Models:	(1)	(2)	(3)
Dependent Variable	$HICP_{NL}$	$HICP_{NL}$	$HICP_{NL}$
$HICP_{EA}$	0.887456***	0.875521***	0.771097***
	{0.017762}	{0.025625}	{0.018406}
	(49.96290)	(34.16698)	(41.89273)
da_{NL}	-0.145186**	-0.034042	
	{0.068497}	{0.041843}	
	(-2.119611)	(-0.813563)	
da_{EA}	0.091386		
	{0.057474}		
	(1.590027)		
l_{NL}	-0.142379**	-0.073385	
	{0.067274}	{0.093166}	
	(-2.116407)	(-0.787671)	
l_{EA}	-0.160410*	-0.188482	
	{0.081325}	{0.115345}	
	(-1.972445)	(-1.634073)	
DUMMY LVL	0.015057***	0.015546***	0.027533***
2012 VAT	{0.002643}	{0.003788}	{0.002643}
	(5.696695)	(4.104276)	(5.696695)
Constant	5.490052***	5.004621***	1.026374***
	{0.875032}	{1.262405}	{0.082444}
	(6.274115)	(3.964353)	(12.44935)
Adj. R^2	0.997310	0.997260	0.994796
Long Run Var.	2.92E-05	6.08E-05	0.000180
SE of reg.	0.003033	0.003061	0.004218
$\sum \varepsilon^2$	0.001232	0.001265	0.002456
EG - ADF tests on the residuals			
Lags (SIC)	1	1	1
Engle Granger τ-stat.	-4.831420*	-4.676252*	-3.783459*
Engle Granger z-stat.[a]	-44.66431**	-41.35915**	-28.00272**

t-stat in (). Standard errors in {}.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

ADF confidence levels were calculated using the MacKinnon et al. (1996); one-sided p-values.

[a] Also known as normalized autocorrelation coefficient.

year-on-year estimation is behaving as we would expect. The error correction term, estimated by the first cointegrating equation has a value of -0.21 and value of -0.18 after removing insignificant covariates, and is always statistically significant at 1% confidence level. This means that in these model 21% and 18% of the disequilibrium is corrected for each month. The second error correction term is also always highly significant, but it's mean estimates are smaller, both being close to -0.08.

This difference in estimates can be most likely attributed to the fact that the first error correction term encompasses the structural factors, while the second one only the inflation. Thus the first error correction can be interpreted as the adjustment toward equilibrium between the inflation and the structural factors, while the second one only between the inflation.

The table 10 shows that the estimates for the error correction term are very high using the first error correction term with the value -2.4 and -2.3. This indicates that the speed of adjustment is simply too quick when we use the error correction from cointegration model between all variables, and thus the error correction term is not appropriate in this case. We can also see from the figure 1 that the annualized rate behaves very erratically. The second error correction term has a value of -0.31 and a value of -0.38 after removing non-significant coefficients. However, in this case the error correction term is only barely significant at the 10% confidence level.

As well as with the year-on-year estimations we can see that the error correction term from cointegrating equation that includes only inflation is much smaller in magnitude. This is again true because in the second case the error correction term can be interpreted only as a speed of adjustment toward equilibrium between the two inflation rates. Meanwhile, in the first case the term should be interpreted as a speed of adjustment toward joint equilibrium between the inflation rates and structural parameters.

Table 9: Summary of the error correction estimations of the relationship between the year-on-year inflation rate in the Netherlands and the rest of the Euro-area.

Overview of ECM Estimations (Year-on-Year Rate). Sample: 02/2002-09/2013.					
Models:	(1)	(2)	(3)	(4)	(5)
Dependent Variable	π_{NL}	π_{NL}	π_{NL}	π_{NL}	π_{NL}
π_{EA}	0.550317***	0.544393***	0.468189***	0.556822***	0.557271***
	[0.378, 0.723]	[0.392, 0.697]	[0.331, 0.605]	[0.400, 0.714]	[0.412, 0.703]
	(6.301246)	(7.071153)	(6.767887)	(7.031498)	(7.584291)
$L^1\pi_{EA}$	-0.413474***	-0.418530***	-0.386909***	-0.418178***	-0.417994***
	[-0.573, -0.254]	[-0.564, -0.274]	[-0.530, -0.244]	[-0.573, -0.264]	[-0.577, -0.253]
	(-5.134500)	(-5.708425)	(-5.336290)	(-5.347410)	(-5.211000)
$L^1\pi_{NL}$	0.897434***	0.902473***	0.927995***	0.896617***	0.896505***
	[0.866, 0.929]	[0.871, 0.934]	[0.889, 0.967]	[0.870, 0.923]	[0.868, 0.926]
	(55.58763)	(56.99472)	(47.25396)	(66.57256)	(61.17568)
Δda_{NL}	-0.015451	-0.004323		-0.022872	-0.023347**
	[-0.047, 0.016]	[-0.037, 0.028]		[-0.056, 0.010]	[-0.041, -0.005]
	(-0.964215)	(-0.263661)		(-1.358452)	(-2.450388)
Δda_{EA}	0.009650	-0.009762		0.015597	0.015852*
	[-0.008, 0.027]	[-0.034, 0.014]		[-0.007, 0.038]	[-0.000, 0.032]
	(1.104940)	(-0.803271)		(1.388764)	(1.932114)
Δl_{NL}	0.011645	0.014213		0.001208	
	[-0.034, 0.057]	[-0.030, 0.059]		[-0.047, 0.049]	
	(0.505870)	(0.632699)		(0.049624)	
Δl_{EA}	-0.050395*	-0.031627		-0.051446	-0.050598***
	[-0.107, 0.007]	[-0.09, 0.033]		[-0.114, 0.011]	[-0.088, -0.014]
	(-1.747820)	(-0.975419)		(-1.626156)	(-2.696315)
Model (1)ϵ_{t-1}		-0.208473***	-0.179720***		
		[-0.312, -0.105]	[-0.267, -0.093]		
		(-3.978105)	(-4.091535)		
Model (3)ϵ_{t-1}				-0.080752***	-0.081322***
				[-0.141, -0.021]	[-0.135, -0.028]
				(-2.654401)	(-2.988447)
Constant	-0.000706	-0.000491	-0.000327	-0.000677	-0.000686
	[-0.002, 0.000]	[-0.002, 0.001]	[-0.001, 0.001]	[-0.002, 0.001]	[-0.002, 0.001]
	(-1.228937)	(-0.637642)	(-0.672861)	(-0.988202)	(-1.143410)
Adj. R^2	0.923579	0.926855	0.925956	0.924794	0.924794
SE of reg.	0.002471	0.002418	0.002432	0.002451	0.002451
$\sum \varepsilon^2$	0.000806	0.000766	0.000799	0.000793	0.000793
Durbin-Wat. stat.	1.829012	1.821397	1.820204	1.829405	1.828969
Wald F-stat.	1012.039	1007.619	878.2251	1713.763	1970.834
Errors:	HAC	HAC	HAC	HAC	HAC

The μ and σ of the dep. var. was during the sample period 0.019878 and 0.008939 respectively.
HAC standard errors & covariances were estimated using Bartlett kernel and Newey-West automatic lags and bandwidth.
t-statistics in (). Lower (L) and Upper (U) 95% confidence intervals are included in [L, U].
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 10: Summary of the error correction estimations of the relationship between the annualized inflation rate in the Netherlands and the rest of the Euro-area.

Overview of ECM Estimations (Annualized Monthly Rate). Sample: 02/2002-09/2013.					
Models: Dependent Variable	(1) π_{NL}	(2) π_{NL}	(3) π_{NL}	(4) π_{NL}	(5) π_{NL}
π_{EA}	0.594144***	0.586909***	0.604448***	0.511687***	0.441640***
	[0.378, 0.810]	[0.425, 0.749]	[0.455, 0.754]	[0.326, 0.697]	[0.262, 0.621]
	(5.445351)	(7.173767)	(7.994861)	(5.462318)	(4.857702)
$L^1\pi_{EA}$	-0.000399	-0.038959		-0.036922	-0.115815*
	[-0.154, 0.153]	[-0.232, 0.154]		[-0.228, 0.154]	[-0.258, 0.027]
	(-0.005156)	(-0.400140)		(-0.382198)	(-1.655373)
$L^1\pi_{NL}$	0.032914	0.031688		0.022409	0.068688**
	[-0.056, 0.122]	[-0.111, 0.174]		[-0.072, 0.117]	[0.002, 0.135]
	(0.733015)	(0.440416)		(0.470321)	(2.026128)
Δda_{NL}	0.001856	0.047184		0.031722	
	[-0.081, 0.085]	[-0.017, 0.112]		[-0.033, 0.097]	
	(0.044163)	(1.449115)		(0.963009)	
Δda_{EA}	-0.102225**	-0.170048***	-0.128112***	-0.115724**	
	[-0.201, -0.003]	[-0.276, -0.065]	[-0.206, -0.050]	[-0.224, -0.007]	
	(-2.040701)	(-3.188938)	(-3.257704)	(-2.114055)	
Δl_{NL}	-0.040675	-0.035428		-0.040788	
	[-0.118, 0.037]	[-0.106, 0.035]		[-0.109, 0.027]	
	(-1.036253)	(-0.993341)		(-1.187086)	
Δl_{EA}	-0.125692	-0.092268	-0.136266**	-0.077570	
	[-0.324, 0.072]	[-0.241, 0.056]	[-0.244, -0.029]	[-0.290, 0.134]	
	(-1.256897)	(-1.231582)	(-2.502442)	(-0.723886)	
Model (1)ϵ_{t-1}		-2.407493***	-2.304869***		
		[-4.198, -0.617]	[-3.808, -0.802]		
		(-2.660295)	(-3.032405)		
Model (3)ϵ_{t-1}				-0.310471	-0.378352*
				[-0.832, 0.211]	[-0.802, 0.046]
				(-1.176875)	(-1.765354)
Constant	0.008140***	0.009034***	0.008544***	0.008271***	0.008738***
	[0.003, 0.014]	[0.005, 0.014]	[0.005, 0.012]	[0.002, 0.015]	[0.003, 0.015]
	(2.860781)	(3.982954)	(4.358653)	(2.636546)	(2.936616)
Adj. R^2	0.141487	0.215659	0.231055	0.131039	0.117593
SE of reg.	0.071582	0.020668	0.020464	0.021754	0.021866
$\sum\varepsilon^2$	0.071582	0.055956	0.056533	0.061993	0.065024
Durbin-Wat. stat.	1.906392	1.968892	1.918297	1.945850	1.956804
Wald F-stat.	8.868397	9.853259	23.42046	17.83819	7.071603
Errors:	HAC	HAC	HAC	HAC	HAC

The μ and σ of the dep. var. was during the sample period 0.019107 and 0.025038 respectively.
HAC standard errors & covariances were estimated using Bartlett kernel and Newey-West automatic lags and bandwidth.
t-statistics in (). Lower (L) and Upper (U) 95% confidence intervals are included in [L, U].
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

The table 11 and 12 show the autocorrelation tests for models 1 to 5 for both year-on-year and monthly annualized rate estimations. No test can reject the null of no autocorrelation at the 1% confidence level. However in the case of 1st, 4th and 5th model in the case of year-on-year rate they can reject the null of no autocorrelation at the 5%, but in the case of 4th and 5th model only by very small margin.

Table 11: Summary of year-on-year serial correlation LM tests.

Breusch-Godfrey Serial Correlation LM Test: year-on-year model (1)			
F-statistic	2.020040	Prob. F(12,120)	0.0279
Obs*R-squared	23.52784	Prob. Chi-Square(12)	0.0236
Breusch-Godfrey Serial Correlation LM Test: year-on-year model (2)			
F-statistic	1.666051	Prob. F(12,119)	0.0829
Obs*R-squared	20.13751	Prob. Chi-Square(12)	0.0645
Breusch-Godfrey Serial Correlation LM Test: year-on-year model (3)			
F-statistic	1.481651	Prob. F(12,123)	0.1398
Obs*R-squared	17.68133	Prob. Chi-Square(12)	0.1257
Breusch-Godfrey Serial Correlation LM Test: year-on-year model (4)			
F-statistic	1.861714	Prob. F(12,119)	0.0460
Obs*R-squared	22.12868	Prob. Chi-Square(12)	0.0361
Breusch-Godfrey Serial Correlation LM Test: year-on-year model (5)			
F-statistic	1.780750	Prob. F(12,120)	0.0588
Obs*R-squared	21.16207	Prob. Chi-Square(12)	0.0481

While the Engle-Granger approach applied in the previous section already yielded interesting result, it also suffers from some drawbacks. For example, the result from the Engle-Granger approach can be sensitive to the normalization applied to the dependent variable. Moreover, Engle-Granger approach does not allow to include lags of dependent variable in the cointegrating vector which is restrictive. Furthermore, residual based tests such as the EG-ADF tend to lack power because it ignores information about the dynamic interaction between the variables. Finally, and most importantly, the Engle-Granger methodology allows us to test only for one cointegrating relationship (Verbeek, 2008).

Table 12: Summary of annualized monthly rate serial correlation LM tests.

Breusch-Godfrey Serial Correlation LM Test: annualized monthly rate model (1)			
F-statistic	1.720288	Prob. F(12,121)	0.0704
Obs*R-squared	20.54967	Prob. Chi-Square(12)	0.0574
Breusch-Godfrey Serial Correlation LM Test: annualized monthly rate model (2)			
F-statistic	1.210881	Prob. F(12,119)	0.2835
Obs*R-squared	15.23456	Prob. Chi-Square(12)	0.2289
Breusch-Godfrey Serial Correlation LM Test: annualized monthly rate model (3)			
F-statistic	1.266857	Prob. F(12,123)	0.2467
Obs*R-squared	15.40004	Prob. Chi-Square(12)	0.2203
Breusch-Godfrey Serial Correlation LM Test: annualized monthly rate model (14)			
F-statistic	1.266857	Prob. F(12,123)	0.2467
Obs*R-squared	15.40004	Prob. Chi-Square(12)	0.2203
Breusch-Godfrey Serial Correlation LM Test: annualized monthly rate model (5)			
F-statistic	1.492840	Prob. F(12,124)	0.1354
Obs*R-squared	17.79870	Prob. Chi-Square(12)	0.1219

4.4.2 Johansen approach

To alleviate above mentioned drawbacks we can use the procedure developed by Johansen (1988), and subsequently improved by Johansen and Juselius (1990), Johansen (1991) and Johansen (2002). The Johansen method allows for testing for more complex cointegrating equations between the variables.

This approach will allow us to estimate the long run relationship with the vector error correction model (VECM), which can added to our ALDR model to facilitate comparison with the previous section. Furthermore, by using this method we don't have to impose strict exogeneity assumption anymore. Even though the Granger causality tests presented before did not provide much evidence that this assumption would not be satisfied, Johansen approach provides further interesting robustness check.

As in the previous section, we will consider two cointegrating models. The first model will include all variables which were also included in the first cointegrating model in the previous section, while the second one will only include price level in the Netherlands and the rest of the Euro-area. For the sake of brevity we will dub the first model 'large'and the second one 'small'.

Before we can start testing for cointegration with the Johansen approach we have to determine the appropriate number of lags. Table 13 presents the optimal number of lags selected by the conventional lag criteria. For the 'large'model all of the criteria select 12 lags, except for AIC which selects 5 lags. In the case of the 'small' model LR, FPE and AIC select 3 lags and SC and HQ only 1 lag. In the subsequent estimations we will let the number of lags to be determined by the majority vote of the information criteria[31].

[31] Although it is certainly true that not all selection criteria are created equal, it is not uncommon to select lags in this simple manner (Verbeek, 2008).

Table 13: Lag selection criteria for Johansen cointegration test and Vector Error Correction. Sample: 01/2002-06/2016.

Lag Order Selection Criteria: Large Model

Lag	LogL	LR	FPE	AIC	SC	HQ
0	2474.955	NA	2.49e-23	-35.02063	-34.89515	-34.96964
1	4077.158	3045.323	5.60e-33	-57.23628	-56.35793	-56.87935
2	4390.788	569.4266	1.09e-34	-61.17429	-59.54306	-60.51142
3	4670.292	483.6822	3.48e-36	-64.62826	-62.24416	-63.65944
4	4842.900	284.0074	5.08e-37	-66.56596	-63.42898	-65.29120
5	4986.125	223.4708	1.13e-37	-68.08688	-64.19702*	-66.50618
6	5005.975	29.28238	1.47e-37	-67.85780	-63.21507	-65.97116
7	5070.856	90.18954	1.02e-37	-68.26746	-62.87186	-66.07488
8	5171.137	130.8626	4.36e-38	-69.17924	-63.03077	-66.68072
9	5226.457	67.48227	3.59e-38	-69.45328	-62.55193	-66.64881
10	5254.813	32.17711	4.45e-38	-69.34486	-61.69063	-66.23445
11	5308.185	56.02195	3.97e-38	-69.59127	-61.18417	-66.17492
12	5406.418	94.74925*	1.94e-38*	-70.47401*	-61.31403	-66.75171*

Lag Order Selection Criteria: Small Model

Lag	LogL	LR	FPE	AIC	SC	HQ
0	763.5664	NA	5.41e-07	-8.753637	-8.717326	-8.738907
1	1608.989	1661.692	3.41e-11	-18.42516	-18.31622*	-18.38097*
2	1614.212	10.14740	3.37e-11	-18.43922	-18.25767	-18.36557
3	1621.090	13.20202*	3.26e-11*	-18.47230*	-18.21812	-18.36919
4	1621.246	0.296096	3.40e-11	-18.42812	-18.10132	-18.29555
5	1626.300	9.468518	3.36e-11	-18.44023	-18.04081	-18.27820
6	1627.112	1.503710	3.49e-11	-18.40359	-17.93155	-18.21210
7	1629.917	5.126299	3.54e-11	-18.38986	-17.84519	-18.16891
8	1632.217	4.150094	3.61e-11	-18.37031	-17.75303	-18.11990
9	1634.242	3.608114	3.69e-11	-18.34761	-17.65770	-18.06774
10	1634.705	0.813444	3.85e-11	-18.30695	-17.54442	-17.99762
11	1638.208	6.079495	3.88e-11	-18.30124	-17.46609	-17.96245
12	1643.521	9.099760	3.82e-11	-18.31633	-17.40856	-17.94808

* indicates lag order selected by the criterion
LR: sequential modified LR test statistic (each test at 5% level)
FPE: Final prediction error
AIC: Akaike information criterion
SC: Schwarz information criterion
HQ: Hannan-Quinn information criterion

We start by running the Johansen unrestricted cointegration test between the inflation, productivity differentials and labour supply in both the Netherlands and the Euro-area, with 12 lags. The results in table 14 show that both Trace and Maximum Eigenvalue tests cannot reject the hypothesis that there are at the most 5 cointegrating vectors. Moreover, the estimated long-run coefficients are not that different from the set up which includes only price levels. Because of this we will not continue our analysis with this specification[32] .

Table 14: Results from the Johansen unrestricted cointegration test between price levels and structural parameters in the NL and EA. Sample: 01/2002 - 09/2013.

Trend assumption: Linear deterministic trend

Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.518473	271.4042	95.75366	0.0000
At most 1 *	0.362006	168.3624	69.81889	0.0000
At most 2 *	0.263071	104.9932	47.85613	0.0000
At most 3 *	0.206773	61.95107	29.79707	0.0000
At most 4 *	0.176253	29.28901	15.49471	0.0002
At most 5	0.013737	1.950278	3.841466	0.1626

Trace test indicates 5 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.518473	103.0418	40.07757	0.0000
At most 1 *	0.362006	63.36913	33.87687	0.0000
At most 2 *	0.263071	43.04215	27.58434	0.0003
At most 3 *	0.206773	32.66206	21.13162	0.0008
At most 4 *	0.176253	27.33873	14.26460	0.0003
At most 5	0.013737	1.950278	3.841466	0.1626

Max-eigenvalue test indicates 5 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

[32] Although in principle there are methods for estimating VEC model in cases such as these, here were are not interested in estimating pure VEC model but rather use the VEC coefficients to add an error correction term to our ALDR model. Because in any cases the estimated cointegrated coefficients are similar this choice should not have any qualitative impact on the results.

Now we will turn our attention to the Johansen unrestricted cointegration test between inflation in the Netherlands and the rest of the Euro-area alone. We will run the Johansen test with 3 lags allowing for linear deterministic trend. The table 15 shows that when we restrict our analysis of cointegration only to relationship between price levels in the Netherlands and rest of the Euro-area, we really do find a single cointegrating relationship as both Trace and Maximum Eigenvalue test statistics cannot reject the null as there is at most one cointegrating relationship between the variables.

Table 15: Results from the Johansen unrestricted cointegration test between price levels in the NL and Euro-area. Sample: 01/2002 - 06/2016.

Trend assumption: Linear deterministic trend
Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.106398	22.03671	15.49471	0.0045
At most 1	0.014054	2.462670	3.841466	0.1166

Trace test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.106398	19.57404	14.26460	0.0066
At most 1	0.014054	2.462670	3.841466	0.1166

Max-eigenvalue test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

The corresponding VEC model of the cointegrating relationship can be seen below in table 16, which shows the normalized coefficients of the cointegrating equation [33]. As we can see in the table the estimations of the coefficient for the price level in the Euro-area is around 0.80 which is very close to the estimation from the Engle-Granger approach. Moreover, this estimate is highly significant at all conventional confidence levels with the t-statistics -28.9148.

Furthermore, the figure 31 and table 51 show that the model is also dynamically stable. While it is true that there is one root that lies on the unit circle, VEC specification imposes $K - r$ roots where K is the number of endogenous variables and

[33] For those who are interested, table 50 shows the full estimations of VEC model.

r is the number of cointegrating relationship. Here recall that in our specification we relaxed the strict exogeneity assumption for the sake of further robustness check. This means that we let the two series be endogenous, for the purpose of this robustness check. Moreover, since Johansen test found one cointegrating relationship we have $K - r = 2 - 1 = 1$. As long as all other roots lie inside the unit circle, model is considered stable (Hamilton, 1994).

Table 16: Estimations of Cointegrating Equation. Sample: 01/2002 - 06/2016

Cointegrating Eq:	CointEq1
$p_{NL,t-1}$	1.000000
$p_{EA,t-1}$	-0.803118 (0.02778) [-28.9148]
Constant	-0.889180

Standard errors in ( ) & t-statistics in [ ]

The coefficients shown above can be directly used as an error correction term, which is also explicitly shown below:

$$\epsilon_{t-1} = p_{NL,t-1} - 0.803118 p_{EA,t-1} - 0.889180 \tag{37}$$

Substituting this error correction term in the same ALDR model which we have used in the previous section, we can estimate further robustness check to our error correction specification. Both estimations for the year-on-year rate and annualized monthly rate can be seen below in table 17. These estimations come very close to those from previous section. In the year-on-year estimations we find that error correction has statistically significant value of -0.11 and -0.10 which is very close to the value we found in the previous section.

In the case of annualized rate we find that the error correction has values -0.58 and -0.60 and that only the former value is statistically significant. Here the coefficient estimates are farther away from those we found in the previous section but the results are similar in a sense that we again found that the error correction term is larger, in its absolute value, in the annualized rate estimations.

Table 17: Summary of the error correction estimations of the relationship between the inflation rate in the NL and the rest of EA.

Overview of Johansen ECM Estimations (Both Year-on-Year and Annualized Monthly Rate). Sample: 02/2002-09/2013.						
	Year-on-Year Models			Annualized Rate Models		
Models:	(1)	(2)	(3)	(1)	(2)	(3)
Dependent Variable	π_{NL}	π_{NL}	π_{NL}	π_{NL}	π_{NL}	π_{NL}
π_{EA}	0.548976***	0.519366***	0.485771***	0.584341***	0.544675***	0.482208***
	[0.376, 0.722]	[0.354, 0.685]	[0.368, 0.604]	[0.387, 0.782]	[0.349, 0.741]	[0.274, 0.691]
	(6.262754)	(6.203843)	(8.139234)	(5.862522)	(5.491240)	(4.575009)
$L^1\pi_{EA}$	-0.381294***	-0.370349***	-0.362415***	-0.050300	-0.119471	-0.193264***
	[-0.517, -0.246]	[-0.510, -0.231]	[-0.493, -0.232]	[-0.132, 0.032]	[-0.269, 0.030]	[-0.335, -0.052
	(-5.563954)	(-5.256798)	(-5.485255)	(-1.215545)	(-1.580732)	(-2.697069)
$L^1\pi_{NL}$	0.862956***	0.877662***	0.887242***	0.048452	0.080788**	0.128307***
	[0.810, 0.920]	[0.846, 0.910]	[0.861, 0.914]	[-0.017, 0.114]	[0.005, 0.157]	[0.046, 0.211]
	(30.06841)	(54.17882)	(66.15791)	(1.459846)	(2.107877)	(3.085090)
Δda_{NL}	-0.028299	-0.030886		-0.008390	-0.000659	
	[-0.068, 0.011]	[-0.074, 0.012]		[-0.082, 0.065]	[-0.067, 0.066]	
	(-1.418196)	(-1.432798)		(-0.224857)	(-0.019507)	
Δda_{EA}	0.015281	0.020292		-0.075575*	-0.067875	
	[-0.008, 0.038]	[-0.007, 0.048]		[-0.157, 0.006]	[-0.158, 0.022]	
	(1.325241)	(1.473109)		(-1.828311)	(-1.496262)	
Δl_{NL}	-0.010891	-0.029350		-0.066739	-0.084970**	
	[-0.063, 0.041]	[-0.085, 0.026]		[-0.147, 0.014]	[-0.164, -0.006]	
	(-0.414270)	(-1.053235)		(-1.639650)	(-2.119237)	
Δl_{EA}	-0.017804	-0.014646	-0.042901**	-0.064778	-0.062507	
	[-0.079, 0.044]	[-0.085, 0.056]	[-0.077, -0.009]	[-0.225, 0.096]	[-0.232, 0.107]	
	(-0.574534)	(-0.412069)	(-2.493721)	(-0.799426)	(-0.730155)	
(Johansen)ϵ_{t-1}		-0.112424**	-0.104964**		-0.577195	-0.602116*
		[-0.201, -0.024]	[-0.188, -0.022]		[-1.324, 0.169]	[-1.279, 0.075]
		(-2.519297)	(-2.514796)		(-1.529588)	(-1.759476)
Dummy M	0.002227***	0.004207***	0.003669***	0.101618***	0.106453***	0.111137
2012 VAT	[0.000, 0.004]	[0.002, 0.007]	[0.002, 0.006]	[0.096, 0.107]	[0.098, 0.115]	[0.104, 0.118]
	(2.361521)	(3.642007)	(3.876708)	(23.75363)	(23.75363)	(31.99992)
Constant	-0.000830	-0.001397*	-0.001053	0.007961***	0.006924*	0.007531**
	[-0.002, 0.000]	[-0.003, 0.000]	[-0.002, 0.000]	[0.003, 0.013]	[0.000, 0.014]	[0.001, 0.014]
	(-1.404993)	(-1.668613)	(-1.654455)	(3.162295)	(1.954303)	(2.270284)
Adj. R^2	0.930484	0.931980	0.932502	0.254949	0.274177	0.266559
F-Stat.	235.2409	214.1352	323.3548	6.988325	6.876053	11.17621
SE of reg.	0.002427	0.002400	0.002391	0.021612	0.021331	0.021443
$\sum \varepsilon^2$	0.000777	0.000755	0.000766	0.061654	0.059608	0.062073
Durbin-Wat. stat.	1.803601	1.785079	1.791714	1.870349	1.883777	1.888018
Wald F-stat.	534.0554	894.5841	1456.432	1439.384	853.6298	892.6367
Errors:	HAC	HAC	HAC	HAC	HAC	HAC

The μ and σ of the year-on-year dep. var. was during the sample period 0.020073 and 0.009204 respectively.
The μ and σ of the annualized monthly rate dep. var. was during the sample period 0.019107 and 0.025038 respectively.
HAC standard errors & covariances were estimated using Bartlett kernel and Newey-West automatic lags and bandwidth.
t-statistics is included in ().
Lower (L) and Upper (U) 95% confidence intervals are included in [L, U].
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

As in previous section, the serial correlation LM tests cannot reject the null hypothesis of no autocorrelation in the residuals from year-on-year estimations at 5% confidence level. For annualized monthly rate, there are few cases where the null hypothesis can be rejected only at 1% confidence interval. Once again, individually such finding could be worrying. However, the fact that the test statistics improves after removing insignificant covariates and that this is not a problem in estimations utilizing longer sample, mitigates these worries.

Table 18: Overview of Berusch-Godfrey Serial correlation LM tests.

Breusch-Godfrey Serial Correlation LM Test: year-on-year model (1)			
F-statistic	1.643395	Prob. F(12,120)	0.0885
Obs*R-squared	19.90130	Prob. Chi-Square(12)	0.0690
Breusch-Godfrey Serial Correlation LM Test: year-on-year model (2)			
F-statistic	1.529040	Prob. F(12,119)	0.1230
Obs*R-squared	18.83629	Prob. Chi-Square(12)	0.0926
Breusch-Godfrey Serial Correlation LM Test: year-on-year model (3)			
F-statistic	1.685861	Prob. F(12,122)	0.0778
Obs*R-squared	20.05533	Prob. Chi-Square(12)	0.0660
Breusch-Godfrey Serial Correlation LM Test: annualized monthly rate model (1)			
F-statistic	1.787620	Prob. F(12,120)	0.0576
Obs*R-squared	21.38298	Prob. Chi-Square(12)	0.0450
Breusch-Godfrey Serial Correlation LM Test: annualized monthly rate model (2)			
F-statistic	2.179827	Prob. F(12,119)	0.0168
Obs*R-squared	25.40865	Prob. Chi-Square(12)	0.0130
Breusch-Godfrey Serial Correlation LM Test: annualized monthly rate model (3)			
F-statistic	1.818625	Prob. F(12,123)	0.0520
Obs*R-squared	21.24734	Prob. Chi-Square(12)	0.0469

4.4.3 Conditional (Unrestricted) ECM

The last robustness check which we will consider here is estimating the error correction term using the conditional or unrestricted error correction model of Pesaran and Shin (1998) and Pesaran et al. (2001). The advantage of this approach is that it is easy to implement as it involves only one equation set-up as opposed to two stage approach which we used in the previous two subsections. Moreover, this approach can be used even with a mixture of I(0) and I(1) data.

Estimating this version of the error correction model will modify the model we have used in previous sections into the following form:

$$\begin{split}\pi_{NL,t} = \alpha + \sum_i \beta_i \pi_{NL,t-1-i} + \sum_i \Lambda_i \pi_{EA,t-i} + \eta \Delta da_{NL,t} + \iota \Delta da_{EA,t} \\ + \omega \Delta l_{NL,t} + \upsilon \Delta l_{NL,t} + \mu p_{NL,t-1} + \varphi p_{EA,t-1} + \tau da_{NL,t-1} + \psi da_{EA,t-1} \\ + \varrho l_{NL,t} + \psi l_{EA,t} + \delta_i z_t + \varepsilon_t\end{split} \tag{38}$$

As we can see this expression is very similar to the previous approaches to error correction model. However here we replaced the lagged residuals (ϵ_{t-1}) from the cointegrating equation, by the lagged levels of the differenced variables. This expression makes it clear why this model became known as 'unrestricted', as we do not put any restrictions on the level variables as opposed in the previous approaches, even though originally it was dubbed 'conditional' by Pesaran et al. (2001).

Testing for cointegration in this single equation setting is very simple as it involves only testing the hypothesis that the coefficient of cointegrating variables are jointly indistinguishable from 0 which implies no cointegration, against the alternative that the variables are in fact cointegrated. That is $H_0 : \mu = \varphi = \tau = \psi = \varrho = \psi = 0$. This hypothesis can be tested by the simple Wald test. Nevertheless, unfortunately the classical test statistics cannot be used in this case because the errors do not have normal distribution anymore. However, Pesaran et al. (2001) tabulated the correct critical values so we can use those. The table 19 shows the correct critical values. The critical values between I(0) and I(1) are in the indeterminate region, and the values above I(1) unambiguously indicate cointegrating relationship.

Moreover, because the model uses mix of both differenced and level variables we will exclude dummies from estimations as using control dummies for both levels and differences caused problems with perfect multicollinearity.

Furthermore the models will be estimated over full sample ranging from 01/2001 - 09/2013 and sub-sample ranging from 01/2002 - 09/2013.

Table 19: Critical Value Bounds. Source: Pesaran et al. (2001).

Significance	I0 Bound	I1 Bound
10%	2.26	3.35
5%	2.62	3.79
2.5%	2.96	4.18
1%	3.41	4.68

Critical values assume unrestricted constant & no trend, k=5.

The estimations of the unrestricted error correction model can be seen on table 20. The models are generally consistent with those from the main results section. Moreover, the level variables have the correct signs in cases when they are significant. The tables in the appendix show the result of the Wald tests for the year-on-year and annualized monthly rate models. We can see that we find cointegration in the full sample unambiguously at all confidence levels. In the restricted sample we find that the test statistics is in the indeterminate region. Nevertheless in the light of results from the full sample there is a good reason to believe that the series are still cointegrated.

The long term coefficients can be extracted by dividing each individual coefficient of the level variables (except for the price level in the NL naturally) by dividing their respective coefficients by the coefficient of the price level in the Netherlands. Thus the long term coefficient of price level in the rest of the Euro-area is $-(\varphi/\mu) = -(0.217759/-0.199367) \approx 1.1$ which means that in the long run 1% increase in the price level in the rest of the Euro-area will lead to expected 1.1% increase in the price level in the Netherlands which is very appealing result as it comes close to what we would expect based on the theory from the first section. To some, this coefficient might seem to contradict the long run coefficients found in the previous section but this is not so. The long run coefficients from this section are not directly comparable since here we had problem with including dummies, and also we could utilise full sample because bound tests do not suffer so heavily from the presence of structural breaks like the Engle-Granger or Johansen approach do.

However, if we redo the estimations of model (1) from the table 8 excluding the tax dummies and over full sample we find long run coefficient of 1.06 which is very close to what we found here, but EG-ADF cannot reject the null of no cointegrating relationship. Repeating the same excercise using Johansen approach we find long run coefficient of 1.11, but here the test cannot reject null of having at most 5 cointegrating relationships, which is again most likely the result of the large structural break in 2001.

Table 20: Summary of the unrestricted error correction estimations of the relationship between the year-on-year inflation rate in the Netherlands and the rest of the Euro-area.

Overview of ECM Estimations (Both Year-on-Year and Annualized Monthly Rate). Sample: Full 01/2001-09/2013, Restricted: 02/2002-09/2013.

	Year-on-Year Models			Annualized Rate Models		
Models:	(1)	(2)	(3)	(1)	(2)	(3)
Dependent Var.	π_{NL}	π_{NL}	π_{NL}	π_{NL}	π_{NL}	π_{NL}
Sample	Full	Restricted	Restricted	Full	Restricted	Restricted
π_{EA}	0.441024***	0.469616***	0.553979***	0.427987***	0.604136***	0.400926***
	[0.277, 0.605]	[0.285, 0.655]	[0.377, 0.731]	[0.192, 0.664]	[0.479, 0.729]	[0.131, 0.671]
	(5.311503)	(5.024801)	(6.204857)	(3.590077)	(9.575453)	(2.930938)
$L^1\pi_{EA}$	-0.228004***	-0.276416***	-0.424223***			
	[-0.358, -0.098]	[-0.425, -0.128]	[-0.568, -0.280]			
	(-3.475917)	(-3.677985)	(-5.827998)			
$L^1\pi_{NL}$	0.754764***	0.800687***	0.899701***	-0.113320	-0.000165	-0.017027
	[0.674, 0.836]	[0.725, 0.876]	[0.864, 0.936]	[-0.304, 0.077]	[-0.118, 0.118]	[-0.168, 0.134]
	(18.47252)	(20.97714)	(49.31182)	(-1.175674)	(-0.002765)	(-0.223021)
$L^2\pi_{NL}$				-0.190875***	-0.157259***	-0.222589**
				[-0.277, -0.105]	[-0.248, -0.067]	[-0.416, -0.029]
				(-4.403387)	(-3.428032)	(-2.271239)
Δda_{NL}	0.058530**	0.051090*	-0.032884	0.038934	0.051128	
	[0.007, 0.111]	[-0.010, 0.112]	[-0.076, 0.010]	[-0.064, 0.142]	[-0.055, 0.158]	
	(2.225834)	(1.669571)	(-1.506047)	(0.748806)	(0.949528)	
Δda_{EA}	-0.088759***	-0.078670**	0.019858	-0.130558*	-0.189593***	
	[-0.135, -0.043]	[-0.139, -0.018]	[-0.010, 0.050]	[-0.064, 0.142]	[-0.327, -0.052]	
	(-3.800524)	(-2.568172)	(1.339551)	(-1.715611)	(-2.727856)	
Δl_{NL}	0.055196	0.017682	-0.036093	-0.078322	-0.002539	
	[-0.018, 0.128]	[-0.051, 0.086]	[-0.089, 0.016]	[-0.213, 0.057]	[-0.106, 0.101]	
	(1.499634)	(0.510006)	(-1.362703)	(-1.147349)	(-0.048421)	
Δl_{EA}	0.011224	0.024080	0.000937	-0.115010**	-0.159895**	
	[-0.038, 0.060]	[-0.074, 0.122]	[-0.075, 0.077]	[-0.208, -0.022]	[-0.314, -0.006]	
	(0.452486)	(0.488300)	(0.024472)	(-2.437761)	(-2.054928)	
$L^1 p_{NL}$	-0.199367***	-0.168251***	-0.147914***	-1.8169584***	-2.357057***	-1.252558***
	[-0.248, -0.150]	[-0.267, -0.070]	[-0.224, -0.072]	[-2.272, -1.362]	[-3.019, -1.695]	[-2.126, -0.379]
	(-8.036728)	(-3.386373)	(-3.838159)	(-7.898405)	(-7.042504)	(-2.829838)
$L^1 p_{EA}$	0.217759***	0.183948***	0.117882***	1.583551***	2.1094321***	1.058251***
	[0.164, 0.272]	[0.084, 0.284]	[0.057, 0.178]	[1.115, 2.052]	[1.529, 2.690]	[0.307, 1.810]
	(8.000676)	(3.632379)	(3.856227)	(6.684333)	(7.196034)	(2.780012)
$L^1 da_{NL}$	-0.152849***	-0.155228***		0.137478	-0.251876	
	[-0.219, -0.087]	[-0.235, -0.075]		[-0.602, 0.877]	[-0.789, 0.285]	
	(-4.576867)	(-3.840464)		(0.367714)	(-0.927685)	
$L^1 da_{EA}$	0.120042***	0.126528***		0.183032	0.381075**	
	[0.065, 0.176]	[0.053, 0.200]		[-0.234, 0.600]	[0.026, 0.736]	
	(4.275786)	(3.407307)		(0.867051)	(2.124409)	
$L^1 l_{NL}$	-0.016420	-0.032393		0.583987	0.417920	
	[-0.080, 0.048]	[-0.095, 0.030]		[-0.183, 1.351]	[-0.133, 0.969]	
	(-0.508107)	(-1.023473)		(1.505380)	(1.500038)	
$L^1 l_{EA}$	-0.159579***	-0.109516**		-1.601282 ***	-1.559778***	
	[-0.215, -0.105]	[-0.196, -0.023]		[-2.392, -0.810]	[-2.428, -0.691]	
	(-5.743884)	(-2.499835)		(-4.002020)	(-3.553302)	
Dummy 2012			0.004952***			0.109500***
VAT			[0.003, 0.007]			[0.098, 0.121]
			(4.541480)			(19.02953)
Constant	3.017523***	2.376904***	0.133984***	20.93225***	22.73598***	0.889883**
	[2.196, 3.840]	[1.013, 3.741]	[0.056, 0.213]	[15.368, 26.497]	[14.011, 31.461]	[0.205, 1.575]
	(7.257602)	(3.448870)	(3.373495)	(7.437751)	(5.156377)	(2.565212)
Adj. R^2	0.959277	0.932391	0.928435	0.439564	0.313238	0.154245
F-Stat.	276.4234	149.5167	181.3301	8.386245	5.911942	6.258511
SE of reg.	0.002415	0.002393	0.002391	0.021985	0.020749	0.046121
$\sum \varepsilon^2$	0.000811	0.000727	0.000738	0.067186	0.054678	0.355233
Wald F-stat.	964.8107	530.0780	1082.363	43.36235	19.47560	216.8424

Estimations use HAC errors. HAC standard errors & covariances were estimated using prewhitening based on Schwartz criterion, Bartlett kernel and Newey-West automatic lags and bandwidth.
t-statistics in (). Lower (L) and Upper (U) 95% confidence intervals are included in [L, U].
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

If we redo the estimations in the table 20 over restricted sample adding 2012 VAT hike dummy and removing the levels of structural parameters, making the estimation directly comparable with the model (3) from table 8 and the VEC model shown in table 38, we find coefficient of 0.8 which is exactly the coefficient that we found previously.

As in the previous error correction models, we can see that the annualized monthly rate models have problems with autocorrelation in some cases. This can be seen from the table 21 below. However, as also seen previously these problems disappear after allowing for lags of dependent variable and excluding insignificant variables.

Table 21: The Berusch-Godfrey serial correlation LM tests for both year-on-year and annualized rate 'unrestricted' error correction models.

Breusch-Godfrey Serial Correlation LM Test: year-on-year model (1)			
F-statistic	1.341485	Prob. F(12,127)	0.2033
Obs*R-squared	17.21176	Prob. Chi-Square(12)	0.1418
Breusch-Godfrey Serial Correlation LM Test: year-on-year model (2)			
F-statistic	1.501184	Prob. F(12,115)	0.1335
Obs*R-squared	19.09573	Prob. Chi-Square(12)	0.0862
Breusch-Godfrey Serial Correlation LM Test: year-on-year model (3)			
F-statistic	1.574436	Prob. F(12,127)	0.1084
Obs*R-squared	19.46419	Prob. Chi-Square(12)	0.0779
Breusch-Godfrey Serial Correlation LM Test: annualized monthly rate model (1)			
F-statistic	1.413838	Prob. F(12,127)	0.0906
Obs*R-squared	50.60112	Prob. Chi-Square(12)	0.0540
Breusch-Godfrey Serial Correlation LM Test: annualized monthly rate model (2)			
F-statistic	2.836535	Prob. F(12,127)	0.0018
Obs*R-squared	32.33935	Prob. Chi-Square(12)	0.0012
Breusch-Godfrey Serial Correlation LM Test: annualized monthly rate model (3)			
F-statistic	0.948073	Prob. F(12,155)	0.5010
Obs*R-squared	11.89815	Prob. Chi-Square(12)	0.4539

4.5 Main Findings

The previous section presented a large number of results, robustness checks and technical details. This might be a bit overwhelming for a reader interested only in the main result. Therefore, this sub-section will briefly summarize the main findings of all previous estimations.

First, recall that we started this research by deriving the equation 28, based on previous works of Ca'Zorzi et al. (2005) and more generally on more commonly used international macro models (See the section 2). This relationship postulates that, assuming the Netherlands can be considered as a small country vis-à-vis the rest of the Euro-area, the price level in the Netherlands will be determined by the price level in the rest of the Euro-area, structural factors in both areas and random shocks.

$$p_{NL,t} = p_{EA,t} - \theta_{EA,t} + \theta_{NL,t} + \epsilon_t \tag{28}$$

However, this relationship cannot be estimated by standard OLS, as the price and structural parameter levels are a unit root process. Therefore, this relationship was estimated inside of an error correction model (See the section 3). This model can be generally described by the equation 31 shown below. The difference between aforementioned approaches lie mostly in the different methodologies for estimating or omitting μ (See the subsections 4.1 and 4.4).

$$\pi_{NL,t} = \alpha + \sum_i \beta_i \pi_{NL,t-1-i} + \sum_i \Lambda_i \pi_{EA,t-i} + \upsilon\vartheta_{NL,t} - \nu\vartheta_{EA,t} - \mu\epsilon_{t-1} + \delta_i z_t + \varepsilon_t \tag{31}$$

The most important results of this estimation (i.e. the β_t and Λ_t) are presented in the table 22 below. The table 22 shows the estimates for the impact of change of inflation rate in the rest of the Euro-area on the Netherlands. As can be seen below, the results are quite robust.

When it comes to year-on-year estimates they are nearly identical, conditional on the sample. In the 2001-2013 sample, year-on-year estimates were approximately 0.452 and 0.441 for the short-run and unrestricted error correction model respectively. The 2002-2013 sample year-on-year estimates were approximately 0.550, 0.544, 0.557 and 0.554 for the short-run, Engle-Granger, Johansen and unrestricted error correction model.

Within each sample these estimates are nearly identical. Between the two samples we can see that there is a small difference in the estimates. In the 2001-2013 sample the coefficients are generally smaller than in the 2002-2013 sample. However, this was expected, as there is no reason to assume that the average impact of inflation rate in the rest of the Euro-area on Dutch inflation rate should be constant. Nevertheless, it is worth pointing out that these different mean estimates are, for the most part, still within 95%[34] confidence intervals of each other. Thus, these differences might have been result of an inherent randomness as well.

Results from the annualized rate estimates seem to be slightly less robust, relative to the year-on-year ones. In the 2001-2013 sample, the mean estimates for annualized rate were approximately 0.532 and 0.428 in short-run and unrestricted error correction model respectively. Annualized estimates from the 2002-2013 sample are 0.594, 0.587, 0.545 and 0.401 for short-run, Engle-Granger, Johansen and unrestricted error correction model.

In this case the models are seemingly less robust than in the year-on-year case. This is especially true for the unrestricted error correction model. However, here it is important to remember that unrestricted error correction model was slightly tweaked to get rid of autocorrelation. More specifically, the annualized rate unrestricted error correction model has different short-run dynamics as it does not include a lag of the annualized rate inflation in the rest of the Euro-area. This could account for the difference in the estimates. Furthermore, as in the previous case, virtually all these estimates are laying within each other's 95% confidence intervals. Therefore, again some of the difference in the mean estimates could be accounted for by sheer randomness.

Comparing the year-on-year estimates with their annualize rate counterparts, we can see that they are still very similar except for the short-run estimates in the 2001-2013 sample and unrestricted error correction model in the 2002-2013 sample. These two pairs of estimates are much more different than all the other ones. However, once again they are for the most part still within 95% confidence intervals of each other.

We can also see that the inflation coefficients are consistently larger in the annualized rate models. This can be due to the fact that annualized rate of inflation varies more. However, it can be a statistical artefact as well, since these results are within 95% confidence intervals of each other, and thus they could be result of a random chance.

[34]These confidence intervals, as well as F-tests and other important statistics, can be seen in the tables containing the full estimation output.

Table 22: Overview of the estimated impact of inflation in the rest of the Euro-area on the Netherlands

π_{EA}:	SR YOY	SR AR	EG YOY	EG AR	JOH YOY	JOH AR	U YOY	U AR
Sample/Model	(1)	(1)					(1)	(1)
2001-2013	0.452***	0.532***					0.441***	0.428***
Sample/Model	(1)	(1)	(2)	(2)	(3)	(3)	(3)	(3)
2002-2013	0.550***	0.594***	0.544***	0.587***	0.557***	0.545***	0.554***	0.401***

The values were rounded from 7 to 3 decimals.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
The values were rounded from 7 to 3 decimals.
SR: Short-Run
EG: Engle-Granger error correction.
Joh: Johansen error correction.
U: Unrestricted error correction.

The table 23 shows an overview of the estimated coefficient of price level in the Netherlands from the estimations of cointegrated (long-run) relationship (see the subsection 4.4). The Engle-Granger and Johansen cointegrated equation was same for both year-on-year and annualized rate estimation, thus they share these coefficients.

In the 2002-2013 sample all price level coefficients are quite similar. This holds less so for the 2001-2013 sample where the coefficients differ bit more between the two versions of unrestricted error correction model. However, it is important to remember that the year-on-year and annualized rate unrestricted error correction model have slightly different specifications. To be more specific the annualized rate version of unrestricted error correction model has different lag specification which was necessary to get rid of autocorrelation.

Contrary to the estimated inflation rate coefficients, the price level coefficients are higher in the 2001-2013 sample than in the 2002-2013 sample. This might be because in a longer sample the long run cointegrated relationship, hidden within error correction term, has larger impact in a longer time span. However, this might again just be due to pure chance as the coefficients are not widely different from each other.

It is also worth pointing out that following the equation 28, we would expect these long run coefficients to be equal to 1. With the exception of 2001-2013 year-on-year version of unrestricted error correction model, estimates are centred more around 0.8. Nevertheless, this difference is still (economically) small and can be a result of some unaccounted differences in taxes or other policies that can affect price levels between the countries[35]. Due to large diversity in taxes and other policies across the Euro-area and numerous changes in these during the sample period, this

[35]SStatistically speaking the standard errors of these coefficients are small, and thus this result

research could correct only for changes in the Dutch VAT rate. Moreover, some of this difference can be attributed also to rough approximation of the structural factors in the Netherlands and the rest of the Euro-area.

Table 23: Overview of the long run relationship between the respective price levels in the rest of the Euro-area and the Netherlands.

π_{EA}:	Engle-Granger	Johansen	Unrest. YOY	Unrest. AR
Sample/Model			(1)	(1)
2001-2013			1.092***	0.872***
Sample/Model	(3)	(2)	(3)	(3)
2002-2013	0.771***	0.803***	0.797***	0.845***

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
The values were rounded from 7 to 3 decimals.

Another important set of results can be seen on table 24 which shows the overview of the estimated error correction term. Because the unrestricted error correction model does not include 'traditional' single error correction term, the net effect of cointegrated level variables was included instead [36].

In the 2001-2013 sample, the error correction of annualized rate version of unrestricted error correction model, the estimate exceeds one. This implies that the error correction is redundant in the given estimation[37], and this is actually result which was also found in some other annualized rate Engle-Granger and Johansen error correction models (but not in the preferred ones which are reported below). This could be due to large volatility in the annualized rate of inflation, as the annualized rate has about twice the variance of the year-on-year inflation rate. Moreover, all annualized rate error correction terms are larger than their year-on-year counterparts. Furthermore, while the year-on-year estimate of -0.174 lies within the boundaries of error correction term, it is also larger than the preferred error correction models from the 2002-2013 sample[38]. Because the error correction coefficient can be interpreted

is unlikely due to just random chance. However, as it was already stressed, it was not possible to take into an account all subtle differences in policies between the Netherlands and the rest of the Euro-area, that can affect their respective prices. An indication of this is that the coefficients get closer to unity when the tax dummies are explicitly included in the model.

[36]This was done by adding up the estimated coefficients of level variables. The significance was determined by the significance of each individual level variable.

[37]In such case, the adjustment toward equilibrium would be greater than 100%. The implication of this is that the adjustment occurs before next time period, making the error correction term redundant.

[38]Although, it is also close to some of the Engle-Granger estimates (see the subsection 4.4.1.)

as a percentage of disequilibrium corrected in each time period, coefficient of -0.174 implies that 90% of disequilibrium will be corrected within 13 months from the shock that lead to disequilibrium, 99% within 25 months.

All error correction terms, from the preferred models, in the 2002-2013 sample are within the standard boundaries for the error correction coefficient[39]. Starting with the year-on-year error correction coefficients, we can see that the Engle-Granger error correction coefficient of -0.081 implies that 90% of disequilibrium will be corrected in 28 months and 99% of disequilibrium in 55 months. The Johansen error correction coefficient of -0.105 suggests that it will take 21 and 42 months to correct 90% and 99% of disequilibrium respectively. Lastly, the unrestricted error correction coefficient of -0.030 indicates that it will take 76 months to correct 90% and 152 to correct 99% of disequilibrium.

Regarding the annualized rate error correction coefficients, we can see that the Engle-Granger coefficient is -0.378. This means that 90% of the disequilibrium will be corrected in 5 months and 99% in 10 months. The Johansen error correction coefficient with value -0.602 implies that the 90% of disequilibrium will be corrected in 3 months and 99% will be corrected in 5 months. Finally, the unrestricted error correction coefficient of -0.194 reveals that 90% and 99% of disequilibrium will be corrected in 11 and 22 months respectively.

There is seemingly a lot of variation between the models in the time that it takes for the disequilibrium to be corrected. However, this is partly due to the nature of exponential decay. Similarly as with compound interest, here too very small differences can easily accumulate over time. Thus, while the aforementioned estimations of time for the disequilibrium correction might be of great interest to some policy makers, they are not best suited for comparisons between the models. To alleviate this drawback the table 24 shows also half-life of decay implied by the error correction. Half-life of decay in the year-on-year models ranges from about 6 to almost 22 months, and in the annualized rate models it ranges from nearly 1 month to 3 months. While there is still considerable variation between these implied half-lives, they all seem quite plausible Also, an important caveat here is that these calculations presuppose that the error correction is constant in each time period, and thus they should be interpreted carefully.

Again it is worth noting that in some cases, these coefficients lie within a 95% confidence interval of each other. Some of these differences could also be explained partially by a random chance. Nevertheless, it has to be admitted that the error correction coefficients are less robust than the inflation rate or price level coefficients.

[39]However, there are few annualized rate models where the error correction coefficient exceeds -1.

Table 24: Overview of the estimated error correction term estimations

π_{EA}:	EG YOY	EG AR	JOH YOY	JOH AR	U YOY	U AR
Sample/Model					(3)	(3)
2001-2013					-0.174***	-1.835***
Implied Half-Life:					3.626	NA
Sample/Model	(5)	(2)	(3)	(3)	(1)	(1)
2002-2013	-0.081***	-0.378*	-0.105**	-0.602*	-0.030***	-0.194***
Implied Half-Life:	8.206	1.460	6.248	0.752	22.757	3.214

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Implied half-life measures the time required for 1/2 of shock to die out. Half-life is defined as:

$t_{1/2} = \frac{tln(2)}{ln(N_0) - ln(N(t))}$, where t is measured in months.

The values were rounded from 7 to 3 decimals.

SR: Short-Run

EG: Engle-Granger error correction.

Joh: Johansen error correction.

U: Unrestricted error correction.

On the other hand, some of these differences could also stem from differences in the way a long-run relationship was modelled. For example, the structural factors had to be excluded from the Johansen error correction model to avoid having multiple cointegrated relationship. But these structural factors were still included in the unconstrained error correction model. Thus, while a lot of effort was exerted to make the models as comparable as possible, small differences that could have some effect on these findings remained.

Other covariates are generally a bit less robust, except for some of the tax dummies. However, these will not be reported in this section. These are not the main focus of this work and interested readers can find them in the previous subsections of this chapter.

5 From Empirics to Forecast

In the previous sections we found that models developed there had a very good fit, without being troubled by endogeneity. This was especially true for the year-on-year models which had very high R^2. This is an essential feature of good forecasting model, and this section will further explore the forecasting abilities of the models developed above. However, this section will be only exploratory. Developing the best possible forecasting model, based on the methodology outlined in the previous sections, is beyond the scope of this work, because such feat would require to extend this work significantly beyond its original purpose. That being said, this chapter might be still of great use to practical forecasters, as it provides a basis on which better models can be built.

In this part we will also introduce several small changes. To start with, here we will use an unemployment rate instead of hours worked and exclude the productivity differentials completely. This had to be done as the data on hours worked is available only up to 2014. Moreover, statistics on hours worked usually lag behind the statistics on unemployment. Thus, while using hours worked is theoretically preferred, doing so would be impractical for forecasting purposes. Furthermore, the unemployment statistics is published on monthly basis and thus we can avoid interpolation which would be impractical for forecasting. Inasmuch, using unemployment is more practical for forecasting purposes, even though hours worked would be preferred from the theoretical perspective. It is true that unemployment is an exact opposite of labour supply, but both are just different sides of the same coin and approximate the labour market responses. The relationship described by equation 28 should also hold for unemployment, albeit with opposite sign.

Additionally, here we will use seasonally unadjusted data as this is again preferable for practical forecasting (Carnot et al., 2011). With this exception the methodology used in the previous chapters will be followed closely.

Moreover, here we will restrict ourselves only to the error correction models based on Johansen (2002) approach to cointegration and the unrestricted error correction model based on Pesaran et al. (2001). The error correction model based on the Engle and Granger (1987) approach will be excluded as it does not offer significant advantages over the previous two[40].

Since in this section we use slightly different variables a new error correction term had to be created with the Johansen method. To do this we will once again have to determine the optimum number of lags. The table 25 reports the results of lag

[40]In fact its main purpose in the previous estimations was to serve as a robustness check to the other methods.

order selection criteria. These tests include the seasonally unadjusted logs of price levels in the Netherlands and the rest of the Euro-area, as well as their respective unemployment numbers.

Table 25: Lag selection criteria for Johansen cointegration test and Vector Error Correction. Sample: 01/1996-06/2016.

Lag Order Selection Criteria

Lag	LogL	LR	FPE	AIC	SC	HQ
0	1089.936	NA	9.31e-10	-9.442924	-9.383131	-9.418804
1	3266.408	4258.314	6.46e-18	-28.22963	-27.93067	-28.10904
2	3360.144	180.1356	3.29e-18	-28.90560	-28.36746	-28.68852
3	3415.109	103.7176	2.34e-18	-29.24443	-28.46712*	-28.93088
4	3438.052	42.49495	2.21e-18	-29.30480	-28.28833	-28.89478
5	3463.474	46.20133	2.03e-18	-29.38673	-28.13109	-28.88023
6	3494.064	54.52961	1.79e-18	-29.51360	-28.01878	-28.91062
7	3529.398	61.75853	1.52e-18	-29.68173	-27.94774	-28.98227
8	3552.924	40.30090	1.43e-18	-29.74717	-27.77401	-28.95124
9	3580.057	45.53574	1.30e-18	-29.84397	-27.63165	-28.95157
10	3598.073	29.60949	1.28e-18	-29.86151	-27.41001	-28.87262
11	3610.764	20.41507	1.33e-18	-29.83273	-27.14206	-28.74737
12	3633.315	35.49279	1.26e-18	-29.88969	-26.95985	-28.70785
13	3709.164	116.7424*	7.55e-19*	-30.41012*	-27.24111	-29.13181*

* indicates lag order selected by the criterion
LR: sequential modified LR test statistic (each test at 5% level)
FPE: Final prediction error
AIC: Akaike information criterion
SC: Schwarz information criterion
HQ: Hannan-Quinn information criterion

Based on the lag selection criteria shown in the table 25 we will use 13 lags for the Johansen cointegration test. The table 26 shows that the trace test founds exactly one cointegrating relationship at the 5% confidence interval. The same holds for the maximum eigenvalue test which also finds only one cointegrating relationship. Since both results are consistent we can conclude that there indeed is a single cointegrating relationship between the price levels and unemployment in the Netherlands and the rest of the Euro-area.

The coefficients of this cointegrating relationship can be seen on the table 27. We can see that the coefficient on inflation has the predicted sign, but its value of ≈ -1.136 is higher then found in previous models. Nevertheless, this is not inconsistent with previous findings. Remember that in those error correction models which took advantages of larger sample also found values of long run coefficient slightly above one. Moreover, here we are using seasonally unadjusted price indices and we also condition on the unemployment.

Unfortunately, it is hard to judge whether the coefficients of unemployment in the

Netherlands and the rest of the Euro-area follow theoretical predictions. The reason for this is that due to lack of data we omit the factor input elasticities from our estimation. These, together with labour supply, enter the relationship multiplicatively ($(b_{iN} - b_{iT})l_i$). Thus the predicted sign will depend on the sign of difference between the factor input elasticity in non-tradable and tradable sectors.

Table 26: Results from the Johansen unrestricted cointegration test between price levels in the Netherlands and the rest of the Euro-area. Sample (adjusted for lags): 03/1997 - 03/2016.

Trend assumption: Linear deterministic trend
Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.120744	54.27755	47.85613	0.0111
At most 1	0.059666	24.80991	29.79707	0.1683
At most 2	0.030633	10.72169	15.49471	0.2292
At most 3	0.015584	3.596930	3.841466	0.0579

Trace test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.120744	29.46764	27.58434	0.0283
At most 1	0.059666	14.08822	21.13162	0.3578
At most 2	0.030633	7.124758	14.26460	0.4745
At most 3	0.015584	3.596930	3.841466	0.0579

Max-eigenvalue test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Using the substituted coefficients we can construct an error correction term, which will take the following form:

$$\epsilon_{t-1} = p_{NL,t-1} - 1.135870 p_{EA,t-1} - 0.084068 l_{NL,t-1} + 0.216268 l_{EA,t-1} - 0.936757 \quad (39)$$

This term will be used to build both year-on-year and annualized monthly rate forecasting model (Although it might also be restimated over a sub-sample for out-of-sample forecasts). These models will be complemented with unrestricted error correction models as a robustness check. Moreover, in contrast to previous chapters the following subsections will be split between year-on-year and annualized monthly rate models, rather than by the technique used for the estimation.

Table 27: Estimations of Cointegrating Equation. Sample (adjusted for lags): 03/1997 - 03/2016.

Cointegrating Eq:	CointEq1
$p_{NL,t-1}$	1.000000
$p_{EA,t-1}$	-1.135870 (0.02589) [-43.8708]
$l_{NL,t-1}$	-0.084068 (0.01627) [-5.16595]
$l_{EA,t-1}$	0.216268 (0.02754) [7.85321]
Constant	-0.936758

Standard errors in ( ) & t-statistics in [ ]

5.1 Year-on-year Forecasting Model

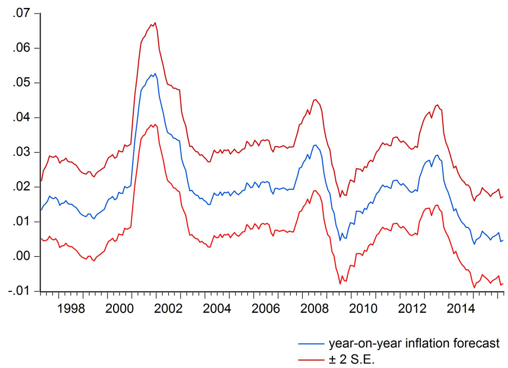
In this section we will look at the forecasting power of the year-on-year models. We will start by estimating the error correction model based on the Johansen approach. The results from these estimations are summarized in table 28.

As the table shows the results are broadly consistent with the models from the previous sections. Moreover, the model still has very a good fit, albeit smaller than before. The error correction is only barely significant, nevertheless this will improve in the subsequent models. Also, this time the model is based on ARDL(3,1) because these are optimal lags according to the Schwartz criterion for the data used in this model. These models do not include year-on-year unemployment, oil or shale revolution dummy, because they were found to be insignificant.

The figure 4a shows the in-sample forecast of the inflation rate in the NL based on the in-sample forecasting model. The model forecast is represented by the blue line, whereas two red lines represent the upper and lower bounds that are 2 standard errors away from the forecast. Additionally, the figure 4b shows the forecasting errors of the model. These were calculated as a difference between the forecast and an actual observed inflation rate.

The figure 4a shows that the forecast generally follows the pattern of the actual realized inflation rate in the Netherlands. More importantly, the figure 4b also shows that the forecasting errors are quite small. For example, between 1997 and 2007 the forecasting errors rarely exceed 0.5% threshold, which is a common threshold used for forecast evaluation (Carnot et al., 2011). The track record of the model is a bit worse after 2007, but for most part the forecast error still does not exceed the 0.5% bounds.

(a) Forecast



(b) Forecast errors

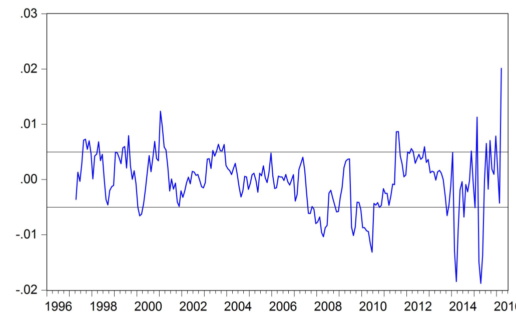


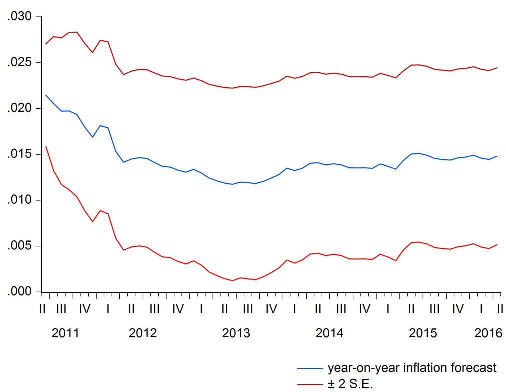
Figure 4: The in-sample forecast and forecast errors of year-on-year inflation rate in the Netherlands from Johansen model.

However, it is known that many models perform quite well when forecasting in-sample, but fail in an out-of-sample forecast. Moreover, it is also important to stress that the in-sample forecasts were estimated as open loop conditional forecast, where independent variables were assumed to follow their actual path. This is of course very unrealistic in a practical setting, exception being situations where this model would be used for a reconstruction of the year-on-year inflation rate. Such assumption will undoubtedly lead to an underestimation of the forecasting error of the model (Carnot et al., 2011).

Therefore, this section will also evaluate the out of sample forecasting power of the model. To do this, we will re-estimate the model over 3/4 of the sample (i.e. 04/1997-06/2011) and use the estimated coefficients to predict inflation in the remaining 1/4 of the sample (i.e. 06/2011-03/2016). Also, one year and one period ahead forecasts will be presented as well. Furthermore, here we will relax our assumptions that the independent variables will follow their actual values. Instead, the out-of-sample values of year-on-year inflation rate are forecasted using simple ARIMA(10,9) for p_{EA}, which was selected by automatic routine based on Schwartz criterion (See the appendix M).

The figure 5a shows the actual forecast of inflation rate and the figure 5b shows the errors from the out of sample forecast. Again, the red lines indicate two standard error away from the forecast. Horizontal lines represent thresholds for errors that exceed 0.5% on both sides. This is a dynamic forecast where the previously forecasted dependent variable is used for further periods. The Euro-area inflation 06/2011-06/2016 is forecast outside the model using ARIMA(10,9), which also incorporated 2001 VAT and 2002 Euro dummies. This auxiliary forecast was based on price levels between 01/1996-06/2016 (See appendix M for further details).

(a) Forecast



(b) Forecast errors

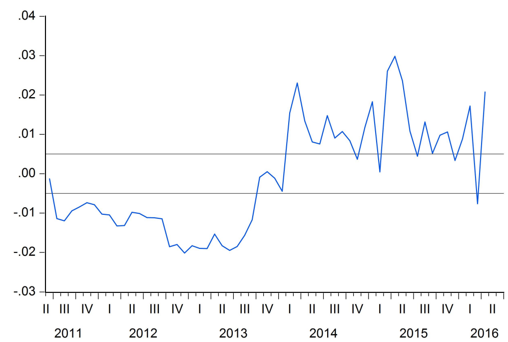


Figure 5: The out-of-sample 4 3/4 year forecast and forecast errors of year-on-year inflation rate in the Netherlands from Johansen model.

As we can see from the figure 5b the model seems to be consistently predicting smaller inflation in the period 2011 to 2013. It also overshoots on average after 2013. This might be due to the fact that the Euro-area was subjected to large volatility almost at all fronts during this period. Also, this was a time when unconventional monetary policy reigned supreme (Fratzscher et al., 2014). In particular, the quantitative easing policy of ECB is not part of this model as it was announced in 2015. Therefore, these conditions put the model through very strict tests.

Using this model for just one step ahead forecasting yields, unsurprisingly, much better results. Now static forecast are naturally useful only for one period ahead forecasting, but such forecasts might still be useful for inflation rate. Especially in the art of practical policy making, knowing inflation rate even one month ahead can be very useful. Moreover, note that these static forecasts still use the estimated inflation rate in the rest of the Euro-area. Thus, these might still be useful in situation where the availability of local inflation rate data is better.

The one step ahead forecasts are shown on figure 6a. These seem to follow the actual inflation rate pretty accurately. Moreover, the standard error bands are quite narrow. The forecasting errors are presented on figure 6b. These lie, for the most part, within 0.5% bounds.

(a) Forecast

(b) Forecast errors

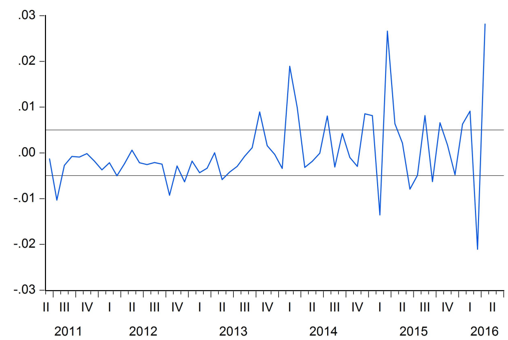


Figure 6: The out-of-sample static forecast and forecast errors of year-on-year inflation rate in the Netherlands from Johansen model.

Next we will continue with examining the one year ahead forecast. To do this we will re-estimate the model once again over the period 04/1997-01/2015. The reason why we estimate new model is to utilize as much data as possible for the estimation of model which will be used for forecasts Moreover, we will also re-estimate the inflation rate in the rest of the Euro-area for the same reason (see appendix M).

The resulting forecast can be seen on the figure 7a, together with bounds based on two standard errors. The figure 7b shows the forecasting errors of this model. The 1 year ahead forecast look better than the 4 3/4 year forecasts. This time most of the forecast errors are within the 5% threshold. Moreover, the 1 year ahead forecast does not seem to be consistently under or overshooting the same way as the 4 3/4 year forecast was.

Next we will turn to the estimation of unrestricted error correction model. The output of this model can be seen in table 29. Again the estimations are generally in line with those we found in the previous sections. The bound tests also find a significant cointegrated relationship (See appendix L). In this case the long run coefficient of price level is about 1.1892 and the joint effect of cointegrated coefficients is -0.00491.

(a) Forecast

(b) Forecast errors

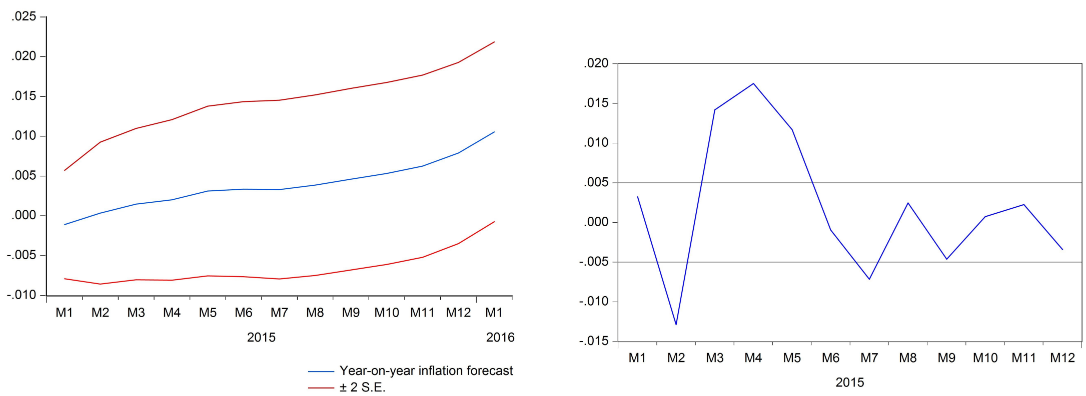


Figure 7: The out-of-sample one year forecast and forecast errors of year-on-year inflation rate in the Netherlands from Johansen model.

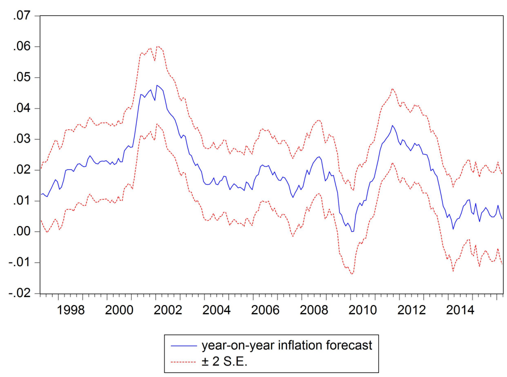
The figure 8a shows forecasting errors of the in-sample unrestricted error correction forecasting model. Again the blue line shows the year-on-year inflation forecast and the red lines show the bounds of 2 standard errors. Furthermore, the figure 8b shows the forecasting errors from this in-sample forecast. Overall, the forecasts look good, but again this were in-sample forecasts conditional on knowing the other independent variables.

Therefore, this section will again also evaluate the out-of-sample performance of the model. This will be done once more by re-estimating the model over 3/4 of the sample, and then producing forecasts for the rest of sample (i.e. 4 3/4 years). Furthermore, we will also present one year (with re-estimated model over 04/1997-01/2015) and one step ahead forecasts as well. These models can be seen in table 29.

The results of the 4 3/4 out-of-sample forecasts from the unrestricted error correction model are shown below on figure 9a. The errors of this forecasting model are shown on the figure 9b. Again the horizontal lines represent the 0.005 or 5% threshold. The forecasting errors of this model seem to generally follow similar patterns of those in the model based on Johansen error correction. However, this time around, the forecasting errors are slightly smaller and the periods of consistent under or overestimating the inflation rate are shorter.

We also estimate the one step ahead forecasts based on this model. This kind of forecast has quite good performance as can be seen from the figures 10a and 10b.

(a) Forecast



(b) Forecast errors

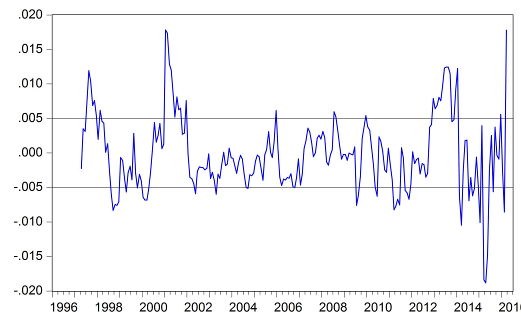
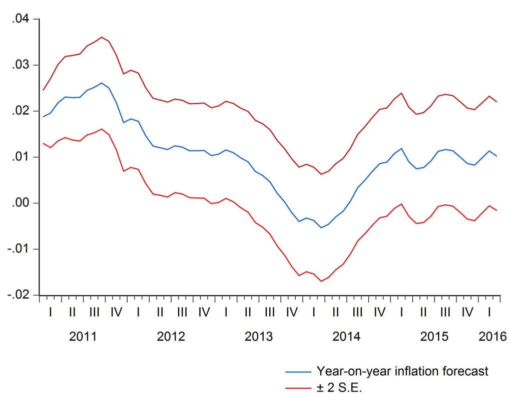


Figure 8: The in-sample forecast and forecast errors of year-on-year inflation rate in the Netherlands from unrestricted error correction model.

(a) Forecast



(b) Forecast errors

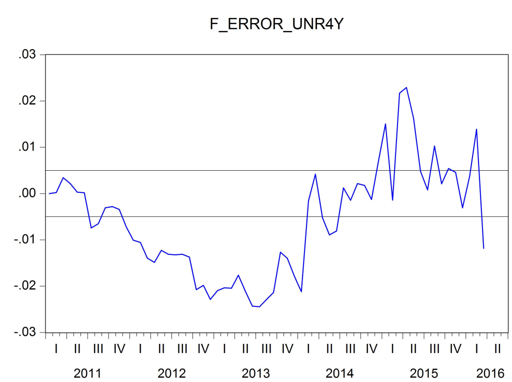
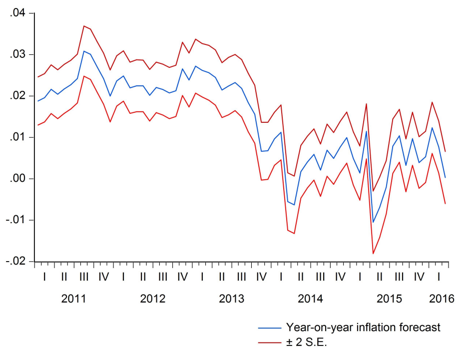


Figure 9: The 43/4 ahead out-of-sample forecast and forecast errors of year-on-year inflation rate in the Netherlands from unrestricted error correction model.

Nevertheless, this kind of forecast has only limited uses and so we will once again continue with estimating model for 1 year ahead forecast.

(a) Forecast



(b) Forecast errors

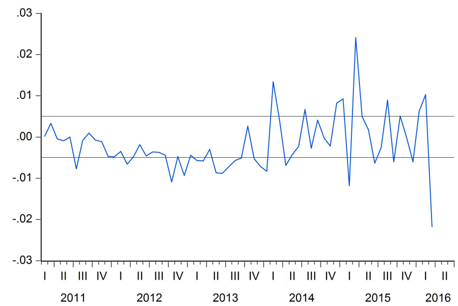
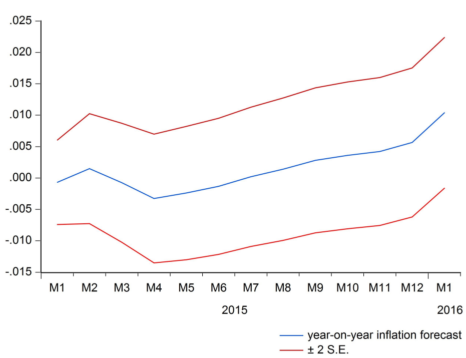


Figure 10: The one step ahead out-of-sample forecast and forecast errors of year-on-year inflation rate in the Netherlands from unrestricted error correction model.

The one year ahead forecast is shown on figure 11a, together with the two standard error bounds. The forecast errors are shown on figure 11b. In terms of forecast errors, this model is the best one from the out-of-sample year-on-year forecasting model.

(a) Forecast



(b) Forecast errors

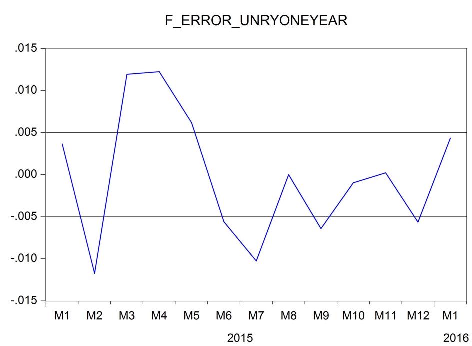


Figure 11: The one year ahead out-of-sample forecast and forecast errors of year-on-year inflation rate in the Netherlands from unrestricted error correction model.

Table 28: Summary of the year-on-year forecasting models based on the Johansen methodology.

Overview of the in-sample and out-of-sample forecasts based on Johansen error correction model. Sample: 04/1997-03/2016.

Models:	In-sample	Out-of-sample	Out-of-sample
Sample:	04/1997-03/2016	04/1997-06/2011	04/1997-01/2015
Dependent Var.	π_{NL}	π_{NL}	π_{NL}
π_{EA}	0.456660***	0.354075***	0.439910***
	[0.223, 0.691]	[0.114, 0.594]	[0.217, 0.663]
	(3.844685)	(2.910867)	(3.893222)
$L^1\pi_{EA}$	-0.314667**	-0.293151**	-0.319023***
	[-0.563, -0.066]	[-0.533, -0.053]	[-0.550,-0.088]
	(-2.493777)	(-2.409209)	(-2.720696)
$L^1\pi_{NL}$	0.661169***	0.823746***	0.854321***
	[0.398, 0.924]	[0.682, 0.965]	[0.735, 0.973]
	(4.953559)	(11.49816)	(14.15910)
$L^2\pi_{NL}$	-0.095005	-0.122315*	-0.238336 ***
	[-0.333, 0.143]	[-0.251, 0.006]	, [-0.418, -0.059]
	(-0.788078)	(-1.876701)	(-2.617241)
$L^3\pi_{NL}$	0.190412**	0.112130**	0.193654***
	[0.069, 0.312]	[0.006, 0.219]	[0.073, 0.314]
	(3.087980)	(2.076287)	(3.173317)
(Johansen)ϵ_{t-1}	-0.029938*	-0.035945**	-0.031860*
	[-0.063, 0.003]	[-0.071, -0.001]	[-0.068, 0.004]
	(-1.772145)	(-2.013247)	(-1.738995)
Dummy	0.010791***	0.008924***	0.009024***
2001 VAT	[0.006, 0.016]	[0.004, 0.014]	[0.005, 0.014]
	(4.321956)	(3.389261)	(3.937117)
Dummy	0.004087**	0.003605**	0.003252**
2002 Euro	[0.001, 0.007]	[0.000, 0.007]	[0.001, 0.006]
	(2.793231)	(2.107590)	(2.382927)
Dummy	0.004681***		0.004014***
2012 VAT	[0.003, 0.007]		[0.002, 0.006]
	(4.424309)		(4.081337)
Constant	0.001423	0.001809*	0.000841
	[0.001, 0.004]	[-0.000, 0.004]	[-0.001, 0.002]
	(1.295750)	(1.745255)	(1.257943)
Forecast Evaluation	In-sample	Out-of-sample	Out-of-sample
Forecast Horizon	whole sample	06/2011-03/2016 ($4\frac{3}{4}$ year)	01/2015-01/2016 1 year
RMSE:	0.006366	0.013716	0.008491
MAE:	0.005024	0.012088	0.006584
Static Fcst. RMSE:		0.008002	
Static Fcst. MAE:		0.005452	
Adj. R^2	0.884204	0.940238	0.923870
SE of reg.	0.004127	0.002740	0.003168
$\sum \varepsilon^2$	0.003713	0.001216	0.002047
Durbin-Wat. stat.	1.741129	1.756184	1.806481
Wald F-stat.	363.0328	497.6410	528.8756
Errors:	HAC	HAC	HAC

HAC standard errors & covariances were estimated using Bartlett kernel and Newey-West automatic lags and bandwidth.
t-statistics included in (). Lower (L) and Upper (U) 95% confidence intervals are included in [L, U].
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
RMSE: Root Mean Squared Error.
MAE: Mean Absolute Error.

Table 29: Summary of the error correction estimations of the relationship between the inflation rate in the Netherlands and the rest of the Euro-area.

Overview of the in-sample and out-of-sample forecasts based on unrestricted error correction model. Sample: 02/2002-09/2013.

Models:	In-sample:	Out-of-sample:	Out-of-sample:
Sample:	04/1997-03/2016	04/1997-06/2011	
Dependent var.	π_{NL}	π_{NL}	π_{NL}
π_{EA}	0.561579***	0.501185***	0.519116***
	[0.24, 0.878]	[0.181, 0.821]	[0.205, 0.834]
	(3.497218)	(3.093546)	(3.255768)
$L^1\pi_{EA}$	-0.273500*	-0.256133***	-0.276760*
	[-0.599, 0.052]	[-0.593, 0.080]	[-0.592, 0.038]
	(-1.654641)	(-1.503705)	(-1.733922)
$L^1\pi_{NL}$	0.603211***	0.764468***	0.817350***
	[0.353, 0.853]	[0.608, 0.921]	[0.688, 0.947]
	(4.769316)	(9.643214)	(12.44792)
$L^2\pi_{NL}$	-0.074307	-0.130273	-0.230158 **
	[-0.283, 0.135]	[-0.292, 0.031]	[-0.410, -0.051]
	(-0.701439)	(-1.595087)	(-2.526529)
$L^3\pi_{NL}$	0.216169 ***	0.079651	0.220638***
	[0.077, 0.356]	[-0.027, 0.186]	[0.090, 0.351]
	(3.053653)	(1.473494)	(3.329530)
Δu_{NL}	-0.017179***	-0.014761***	-0.013059***
	[-0.025, -0.009]	[-0.022, -0.008]	[-0.019, -0.007]
	(-4.234359)	(-4.173444)	(-4.099377)
Δu_{EA}	0.003769	0.009853**	0.004538
	[-0.008, 0.015]	[0.001, 0.019]	[-0.005, 0.014]
	(0.643947)	(2.175333)	(0.941519)
L^1p_{NL}	0.153912***	0.119776***	0.103714***
	[0.070, 0.238]	[0.042, 0.198]	[0.044, 0.164]
	(3.595935)	(3.026191)	(3.422106)
L^1p_{EA}	-0.183026***	-0.149414***	-0.126470***
	[-0.283, -0.083]	[-0.243, -0.056]	[-0.196, -0.057]
	(-3.593114)	(-3.151313)	(-3.573743)
L^1u_{NL}	-0.015642***	-0.014207***	-0.011716***
	[0.070, 0.238]	[-0.024, -0.005]	[-0.019, -0.004]
	(3.595935)	(-2.930275)	(-3.133359)
L^1u_{EA}	0.039846***	0.026061***	0.029977 ***
	[0.021, 0.059]	[0.010, 0.042]	[0.016, 0.045]
	(4.085500)	(3.279882)	(4.075332)
Constant	-0.153419***	-0.028602	-0.112512***
	[-0.223, -0.084]	[-0.087, 0.030]	[-0.169, -0.056]
	(-4.355437)	(-0.960674)	(-3.933270)
Forecast Evaluation	In-sample	Out-of-sample	Out-of-sample
Forecast Horizon	whole sample	06/2011-03/2016 ($4\frac{3}{4}$ year)	01/2015-01/2016 1 year
RMSE:	0.005520	0.012991	0.007396
MAE:	0.004153	0.010400	0.006094
Static Fcst. RMSE:		0.007012	
Static Fcst. MAE:		0.005426	
Adj. R^2	0.888199	0.937205	0.922745
SE of reg.	0.004055	0.002808	0.003191
$\sum \varepsilon^2$	0.003552	0.001254	0.002057
Durbin-Wat. stat.	1.715814	1.816670	1.877197
Wald F-stat.	171.1454	285.8254	290.1826
Errors:	HAC	HAC	HAC

HAC standard errors & covariances were estimated using Bartlett kernel and Newey-West automatic lags and bandwidth.

t-statistics in (). Lower (L) and Upper (U) 95% confidence intervals are included in [L, U].

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

RMSE: Root Mean Squared Error.

MAE: Mean Absolute Error.

5.2 Annualized Monthly Rate Forecasting Model

After looking at the year-on-year models, this section continues with presentation of annualized rate forecasting models. Here we will estimate the Johansen error correction model and the unrestricted error correction model using annualized inflation rate instead of year-on-year rate. Moreover, since these models are created for forecasting purposes, seasonally unadjusted inflation rate is used. To capture a potential seasonal pattern we will use month dummies in these estimations. The seasonal dummies are added for every month except for January, which will be considered as base month. As before, the coefficients were estimated over the full sample for in-sample forecast and over the 3/4 sample and the 04/1997-01/2015 sample for out-of-sample and one step ahead forecasts.

The output from Johansen error correction models can be seen in the table 30. In this case the inflation coefficient is still more or less robust comparing to the main estimations but other coefficients, including the error correction term, are not. This could be because of the different way of addressing seasonality in this case. However, it could also be a result of using slightly different error correction term which included an unemployment in the Netherlands and the rest of the Euro-area.

The in-sample forecast together with a bounds of two standard errors can be seen on the figure 12a and the forecasting errors are shown on the figure 12b. This time around the forecasting errors are undoubtedly larger than in the previous case. However, it is worth noting that the annualized rate of inflation also varies more violently than the year-on-year rate. What is more worrying is that the forecasting errors seem to get very large in the second part of the sample. Nevertheless, that was something to be expected given that in this specification the error correction model is not significant.

Next, the figure 13a and 13b show the out-of-sample 4 3/4 year ahead forecast of annualized rate inflation using Johansen ECM and its forecast errors. Again the out-of-sample forecasts also use auxiliary forecasts of the Euro-area price levels. In this case the forecasting errors are also quite large. Moreover, the errors also increase with the time as previously.

The same also holds for the one step ahead forecast, shown on figure 14a. The one step ahead forecasts naturally tend to perform much better than models which try to forecast over larger horizon (Carnot et al., 2011). However, as it is plain from the forecast errors shown on the figure 14b, this time the one step ahead forecast performs no better than the out-of-sample 4 and 4/3 year ahead forecast.

The same pattern also holds for the year ahead forecast. These forecasts are shown on the figure 15a and the errors from these forecasts are shown on 15b. These forecasts, unfortunately, seem to be a bit worse than the forecasts from the 4 and

(a) Forecast

(b) Forecast errors

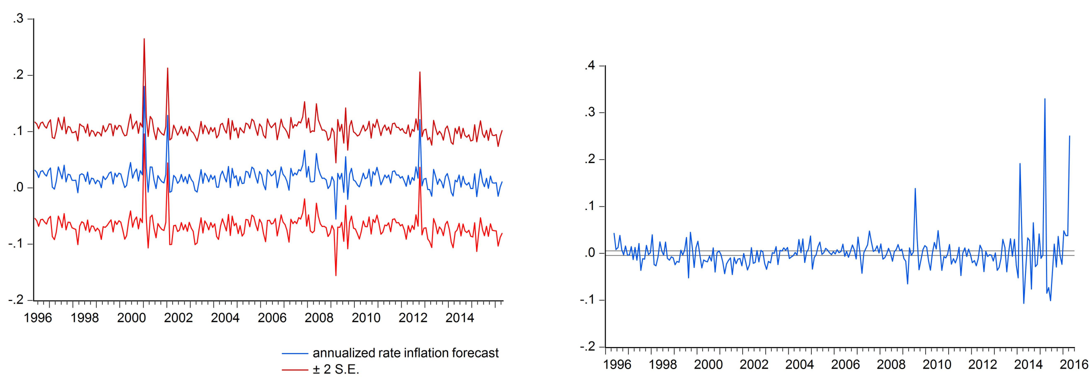


Figure 12: The in-sample forecasts and forecast errors of annualized inflation rate in the Netherlands using Johansen ECM.

(b) Forecast errors

(a) Forecast

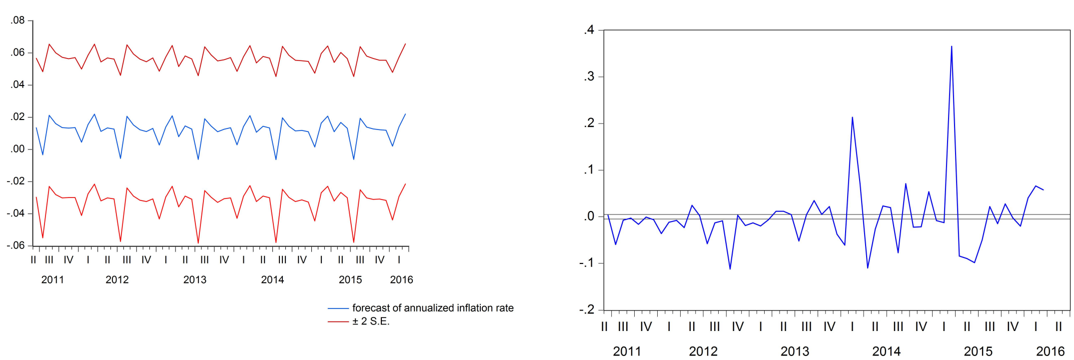


Figure 13: The out-of-sample 4 3/4 year ahead forecasts and forecast errors of annualized inflation rate in the Netherlands using Johansen ECM.

(a) Forecast

(b) Forecast errors

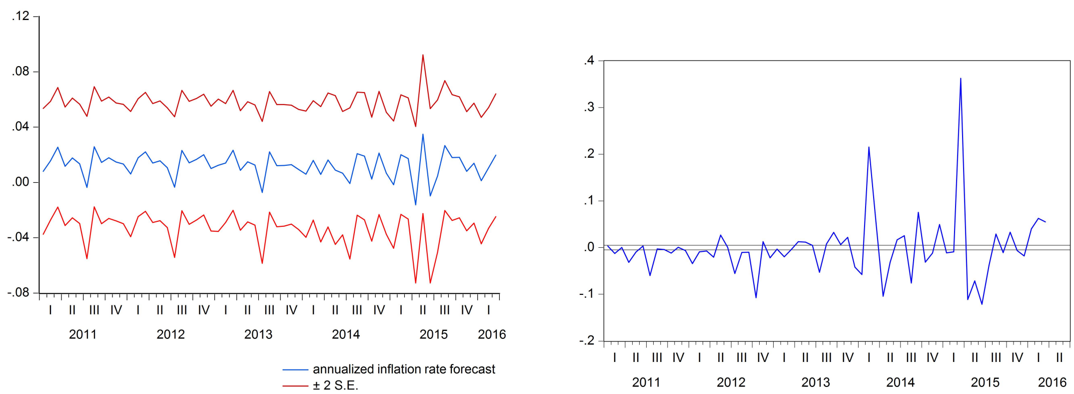


Figure 14: The out-of-sample one step ahead forecasts and forecast errors of annualized inflation rate in the Netherlands using Johansen ECM.

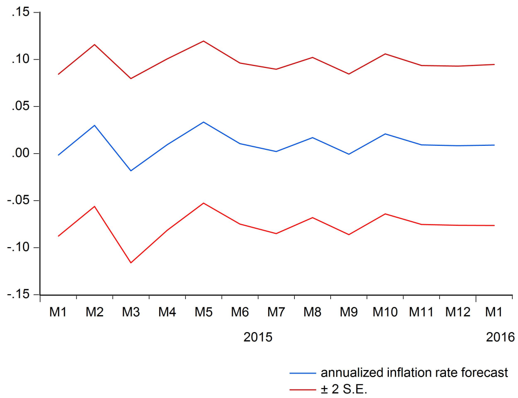
3/4 year ahead forecasts judging by the forecasting errors. This runs to the contrary of what would be expected. Usually the shorter the forecast horizon is, the better the forecasts get. This holds for both Johansen and unrestricted error correction year-on-year models, but paradoxically the annualized rate models perform worse in short sample. Furthermore, the annualized rate forecasting models have generally much larger forecasting errors. Nevertheless, they also vary more and this could explain this finding.

After going through annualized rate versions of Johansen error correction model we will turn our attention to the unrestricted error correction models. Similar to the Johansen error correction models, we find that while the inflation rate estimates are more or less consistent with empirical estimations from previous chapter, other findings are less robust. In this case too error correction mechanics is not significant as the bound tests do not find that error correction terms are significant together[41]. Moreover, unemployment rate was removed from this model as it was insignificant and it inflated errors for other variables as well. Despite of these shortcomings this section also presents forecasts from this model.

The figure 16a and 16b show the in-sample forecast and forecasting errors respectively. We can see that this model performs similarly to the annualized rate Johansen

[41]see appendix L. Also note that while the Wald test is significant at 5%, bound tests do not follow standard asymptotic. Comparing the test statistics to critical values from Pesaran et al. (2001) we find that the error correction terms are not jointly significant at 5%.

(a) Forecast



(b) Forecast errors

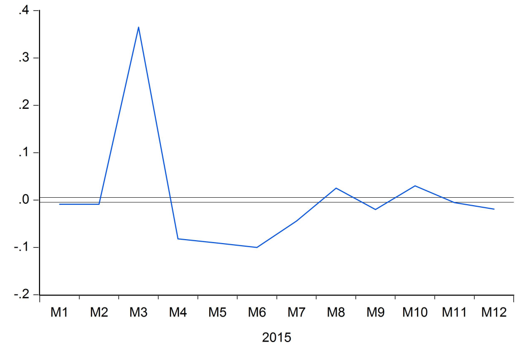
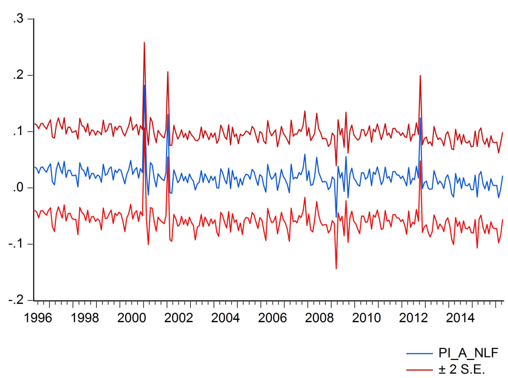


Figure 15: The out-of-sample one year ahead forecasts and forecast errors of annualized inflation rate in the Netherlands using Johansen ECM.

error correction model. This is not surprising as in both models the main difference was the approach to the error correction, but both error correction terms turned out to be not significant. Therefore, the annualized rate Johansen and unrestricted error correction models in this case are virtually same.

(a) Forecast



(b) Forecast errors

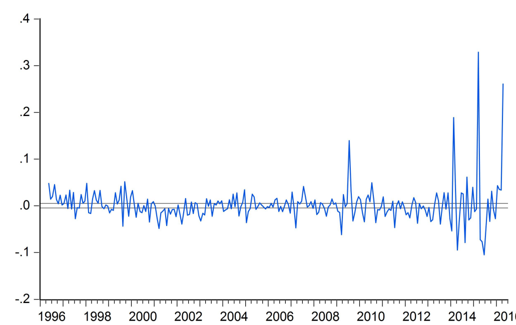
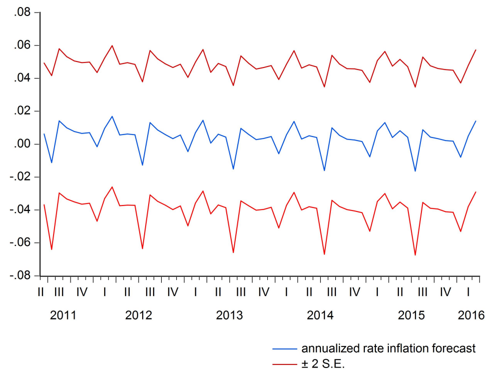


Figure 16: The in-sample forecasts and forecast errors of annualized inflation rate in the Netherlands using unrestricted ECM.

What holds for the in-sample annualized rate unrestricted error correction models, holds for the out-of-sample models as well. The figure 17a, 18a and 19a show the 4 and 3/4 year, one step and one year ahead forecasts respectively. The errors from these forecasts are shown on the figure 17b, 18b and 19b. Again these are almost identical to the ones based on Johansen model.

(a) Forecast



(b) Forecast errors

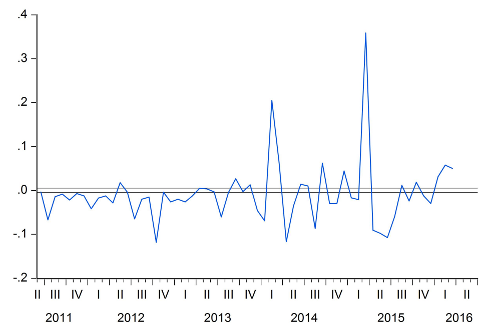
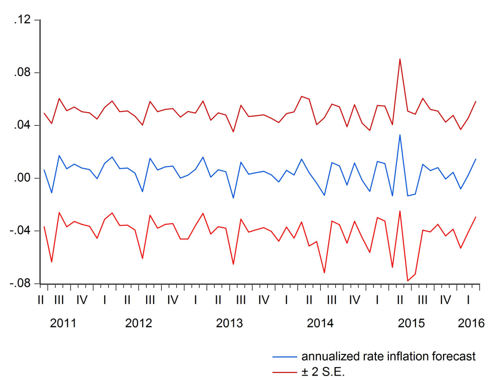


Figure 17: The out-of-sample 4 and 3/4 year forecasts and forecast errors of annualized inflation rate in the Netherlands using unrestricted ECM.

(a) Forecast



(b) Forecast errors

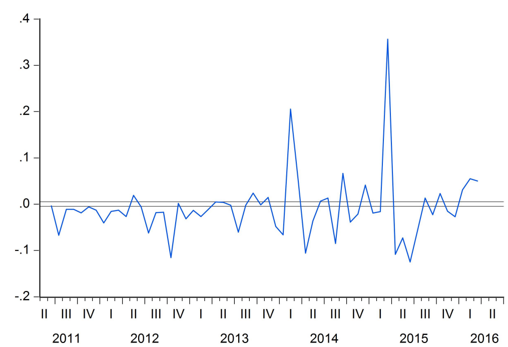


Figure 18: The out-of-sample one step ahead forecasts and forecast errors of annualized inflation rate in the Netherlands using unrestricted ECM.

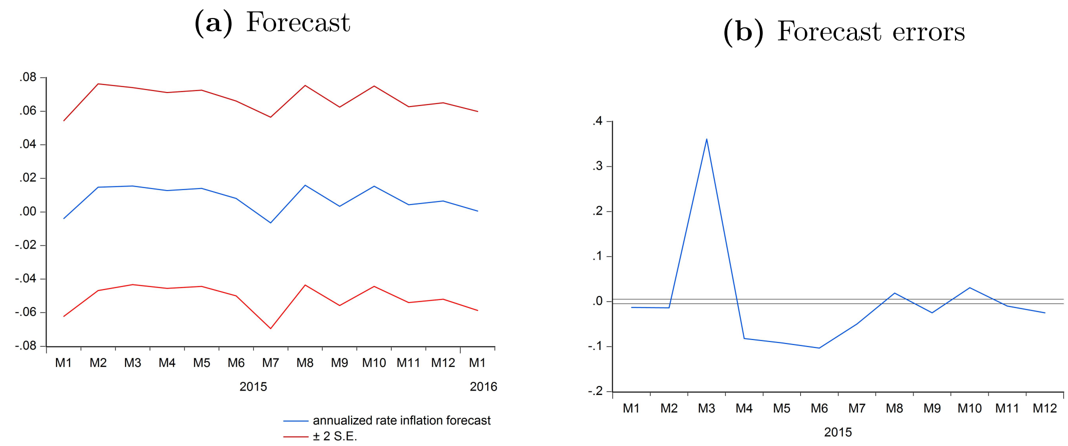


Figure 19: The out-of-sample one year ahead forecasts and forecast errors of annualized inflation rate in the Netherlands using unrestricted ECM.

Table 30: Summary of the annualized rate forecasting models based on the Johansen methodology.

Overview of the in-sample and out-of-sample forecasts based on Johansen error correction model. Sample: 04/1997-03/2016.

Models:	In-sample	Out-of-sample	Out-of-sample
Sample:	04/1997-03/2016	04/1997-06/2011	04/1997-01/2015
Dependent Var.	π_{NL}	π_{NL}	π_{NL}
π_{EA}	0.632898***	0.473909***	0.588796***
	[0.390, 0.876]	[0.279, 0.669]	[0.332, 0.846]
	(5.142652)	(4.803160)	(4.516344)
$L^1\pi_{EA}$	-0.227569*	-0.150639	-0.146989*
	[-0.458, 0.003]	[-0.279, -0.022]	[-0.322, 0.028]
	(-1.948383)	(-2.318575)	(-1.660862)
$L^1\pi_{NL}$	0.009108	0.074742	0.027385
	[-0.179, 0.197]	[-0.025, 0.175]	[-0.113, 0.167]
	(0.095655)	(1.478841)	(0.386107)
$L^2\pi_{NL}$	-0.077665	-0.030127	0.007252*
	[-0.286, 0.131]	[-0.138, 0.077]	, [-0.414, 0.007]
	(-0.734639)	(-0.553823)	(-1.904405)
$L^3\pi_{NL}$	0.063297	0.073810	0.080697
	[-0.145, 0.277]	[-0.053, 0.200]	[-0.033, 0.194]
	(0.585101)	(1.152170)	(1.405670)
(Johansen)ϵ_{t-1}	-0.157771	-0.044180	-0.048092
	[-0.418, 0.102]	[-0.207, 0.119]	[-0.223, 0.127]
	(-1.197655)	(-0.536068)	(-0.542982)
Dummy	0.223570***	0.210676***	0.218106***
2001 VAT	[0.200, 0.247]	[0.185, 0.237]	[0.193, 0.243]
	(18.88837)	(15.97708)	(17.38609)
Dummy	0.103347***	0.102447**	0.096536***
2002 Euro	[0.084, 0.123]	[0.087, 0.118]	[0.081, 0.112]
	(10.64789)	(12.75355)	(12.57554)
Dummy	0.106517***		0.100550***
2012 VAT	[0.090, 0.123]		[0.090, 0.005]
	(12.52883)		(18.59895)
Constant	0.009433	0.007533	0.014545**
	[-0.009, 0.028]	[-0.008, 0.023]	[0.002, 0.027]
	(1.011989)	(0.960118)	(2.268853)
Forecast Evaluation	In-sample	Out-of-sample	Out-of-sample
Forecast Horizon	whole sample	06/2011-03/2016 ($4\frac{3}{4}$ year)	01/2015-01/2016 1 year
RMSE:	0.039457	0.070077	0.112609
MAE:	0.021702	0.040748	0.061564
Static Fcst. RMSE:		0.067444	
Static Fcst. MAE:		0.038255	
Adj. R^2	0.155396	0.379814	0.305422
SE of reg.	0.041460	0.021098	0.027272
$\sum \varepsilon^2$	0.376447	0.067215	0.143548
Durbin-Wat. stat.	1.829321	2.022087	2.109402
Wald F-stat.	5.660987	6.479546	5.683054
Errors:	HAC	HAC	HAC
Month dummies:	Yes	Yes	Yes

HAC standard errors & covariances were estimated using Bartlett kernel and Newey-West automatic lags and bandwidth.
t-statistics in (). Lower (L) and Upper (U) 95% confidence intervals are included in [L, U].
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
RMSE: Root Mean Squared Error.
MAE: Mean Absolute Error.

Table 31: Summary of the error correction estimations of the relationship between the inflation rate in the Netherlands and the Euro-area.

Overview of the in-sample and out-of-sample forecasts based on unrestricted error correction model. Sample: 02/2002-09/2013.

Models:	In-sample:	Out-of-sample:	Out-of-sample:
Sample:	04/1997-03/2016	04/1997-06/2011	04/1997-01/2015
Dependent var.	π_{NL}	π_{NL}	π_{NL}
π_{EA}	0.615264***	0.512620***	0.600397***
	[0.404, 0.827]	[0.353, 0.673]	[0.339, 0.861]
	(5.729401)	(6.332715)	(4.535799)
$L^1\pi_{EA}$	-0.178303	-0.102761 **	-0.105485
	[-0.417, 0.060]	[-0.200, -0.006]	[-0.272, 0.061]
	(-1.473197)	(-2.097006)	(-1.247874)
$L^1\pi_{NL}$	-0.037787	0.048095	0.022655
	[-0.178, 0.103]	[-0.043, 0.139]	[-0.123, 0.169]
	(-0.530180)	(1.048720)	(0.306268)
$L^2\pi_{NL}$	-0.135474	-0.053853	-0.210751**
	[-0.307, 0.036]	[-0.174, 0.067]	[-0.383, -0.039]
	(-1.557370)	(-0.883491)	(-2.412831)
$L^3\pi_{NL}$	-0.006662	0.048525	0.088286
	[-0.133, 0.120]	[-0.095, 0.192]	[-0.033, 0.209]
	(-0.103746)	(0.670007)	(1.438427)
L^1p_{NL}	-0.237742**	-0.085433	-0.183686**
	[-0.471, -0.005]	[-0.208, 0.037]	[-0.336, -0.032]
	(-2.011672)	(0.044865)	(-2.381840)
L^1p_{EA}	0.208733*	0.044865	0.168211**
	[-0.040, 0.457]	[-0.097, 0.186]	[0.008, 0.328]
	(1.656549)	(0.627106)	(2.075341)
Constant	0.140756	0.186520*	0.080751
	[-0.068, 0.349]	[-0.001, 0.374]	[-0.120, 0.281]
	(1.332388)	(1.968372)	(0.795252)
Forecast Evaluation	In-sample	Out-of-sample	Out-of-sample
Forecast Horizon	whole sample	06/2011-03/2016 ($4\frac{3}{4}$ year)	01/2015-01/2016 1 year
RMSE:	0.039476	0.070529	0.111968
MAE	0.021540	0.042822	0.066434
Static Fcst. RMSE:		0.007008	
Static Fcst. MAE:		0.042054	
Adj. R^2	0.198433	0.400818	0.278430
SE of reg.	0.037449	0.020738	0.027797
$\sum \varepsilon^2$	0.288905	0.064509	0.149126
Durbin-Wat. stat.	2.038610	2.037728	2.099806
Wald F-stat.	3.675972	6.685996	5.109477
Errors:	HAC	HAC	HAC
Month dummies:	Yes	Yes	Yes

HAC standard errors & covariances were estimated using Bartlett kernel and Newey-West automatic lags and bandwidth.

t-statistics in (). Lower (L) and Upper (U) 95% confidence intervals are included in [L, U].

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

RMSE: Root Mean Squared Error.

MAE: Mean Absolute Error.

5.3 Comparison of the Forecast Accuracy

This section will compare the results from the previous sections. The table 32 provides the overview of the root-mean-square error (RMSE) and mean squared error (MAE) from both year-on-year and annualized inflation rate forecasts, made by both Johansen and unrestricted error correction model. We will focus mostly on the RMSE as this is standard measure for forecast evaluation. The RMSE represents the sample standard error of estimation, and as a such it is a nice and simple measure of accuracy[42]

Overall we can see that the RMSE of out-of-sample year-on-year models range from about 0.7% for unrestricted ECM one step to around 1.3% for Johansen 4 and 3/4 year ahead forecast. The one year ahead forecast (which is widely used in literature) has RMSE of approximately 0.74%.

These RMSE measures seem to be quite reasonable. For example, Öller and Barot (2000) find that overall inflation models used to forecast inflation among 13 European countries, focusing mainly on core EU countries, is about 1.6%. It is true that for inflation in the Netherlands they report RMSE of 0.39%, but they also focus mostly only on year ahead forecasts. This paper focuses exclusively on a dynamic forecast. Moreover, the forecasts analysed in Öller and Barot (2000) come mostly from 70s to late 90s or early 2000s, which was more stable period.

Indeed, there are scholars who observed that, over time inflation got harder to forecast. Some observers claim this is due to different policies (see Estrella, 2005, for further discussion). Others argue that this can be explained in the context of 'Great Moderation', including changes in the structure of the real economy, the deepening of financial markets, and possible changes in the nature of the structural shocks hitting the economy" (see Stock and Watson, 2007). Also, while these authors focus on the US inflation it is not hard to believe that many of these findings and observations apply to the Euro zone as well.

However, at the same time the performance of the annualized rate forecasting model is looking worse. The RMSE of annualized rate models is quite large, ranging from around 7% to little above 11%. Now here it should be noted that the RMSE can be sensitive to the transformations applied to the dependent variable Hyndman and Koehler (2006), and thus year-on-year and annualized inflation rate models cannot be easily compared against each other. For example, as we noted earlier the annualized rate of inflation has almost two times higher variance than year-on-year

[42]Although this measure is scale dependent. Thus it should be used only for comparing models with the same dependent variable, or variables which are rescaled to be comparable (see Hyndman and Koehler, 2006).

rate. Nevertheless, the annualized rate models seem to perform way worse than could be explained just by this. Moreover, the error correction terms in the annualized inflation rate forecasting model even turned out to be no longer significant.

One plausible explanation for these differences could be a difference in seasonal adjustment between estimations from section 4 and 5. Year-on-year models correct for seasonality by using the year-on-year rates, thus it makes sense that they will perform same as the models in the previous section. However, the annualized rate models were first adjusted for seasonality using TRAMO/SEATS procedure. Such procedure works for the estimation of the relationship, but forecasting often requires use of raw unadjusted data (Carnot et al., 2011). Therefore, the forecasting annualized rate models were rather estimated together with seasonal dummies. Already the annualized rate estimations in chapter 4 were usually less clear, possibly due to large variability of the annualized rate inflation, and inclusion of unadjusted data might have introduced too much noise to the series.

However, all these forecast should be interpreted carefully. As it was mentioned at the beginning of this chapter all these forecasts are exploratory and should be taken as such and there is still room for improvement. For example, as this is multivariate forecasting model, auxiliary forecasts of the rest of the Euro-area inflation rate and Dutch and Euro-area unemployment rates had to be made for the out-of-sample forecasts. These were just created using simple ARIMA models. If models that were explored in the previous section were 'fed' better forecasts of these variables they could perform much better. For example, nowadays unemployment is usually estimated using labour force flows (Barnichon and Nekarda, 2012). Nevertheless, creating more complicated models for these auxiliary forecasts would be outside the scope of this work.

Furthermore, some scholars take the out-of-sample forecasting performance as a test of the underlying theory or estimates. Clements and Hendry (2005) argue that the "out-of-sample forecast performance is not a reliable indicator of the validity of an empirical model, nor therefore of the economic theory on which the model is based." The reason for this is that even many seemingly good forecasting models turn out to be wrong later as the parameters in the economy change and vice versa. Therefore, the relative good out-of-sample performance of the year-on-year forecasting vis-à-vis that of the annualized rate model should not be used too hastily for making judgements about theory and estimates presented in the previous sections.

Table 32: Overview of forecast evaluations.

Model	4 3/4 year forecast	1 step ahead forecast	1 year forecast
RMSE			
Year-on-Year Joh.	1.3716%	0.8002%	0.8059%
Year-on-Year Un.	1.2991%	0.7012%	0.7396%
Annualized rate Joh.	7.007%	6.7444%	11.2609%
Annualized rate Un.	7.0529%	7.0008%	11.1968%
MAE			
Year-on-Year Joh.	1.2088%	0.5452%	0.6584%
Year-on-Year Un.	1.0400%	0.5426%	0.6094%
Annualized rate Joh.	4.0748%	3.8255%	6.1564%
Annualized rate Un.	4.282%	4.2054%	6.6434%

Joh.: Johansen error correction model.

Un.: unrestricted error correction model.

RMSE: Root Mean Squared Error.

MAE: Mean Absolute Error.

6 Possible Extensions

There are several ways this paper could be further extended. First, this paper could be greatly extended by using better dichotomy between tradable and non-tradable sectors. Currently it is commonplace to distinguish tradable and non-tradable sectors by NACE code using the convention set up by AMECO. However, this is very unfortunate as it forces whole sectors into one or other category while reality is more nuanced. It would be better to distinguish sectors using the share of value added embodied in foreign demand (Zeugner, 2013). However, the OECD TiVA database contains this measure only for few years. Additionally, for the purposes of this research we would also need to know the share of labour supply which is devoted to fulfilling the foreign demand. Unfortunately, no such database exists as of yet to the best knowledge of this author. Nevertheless, this work could be greatly improved if such database would come to be.

Second, this paper had to omit shares of income spent on tradable goods as well as the difference between factor input elasticities. These were again omitted due to lack of proper data. However, once again these can be estimated and thus the lack of data could be alleviated in the future. Including these in the analysis would improve precision of the model as well as enable us to better evaluate the underling theory.

Third, this research was applied only to relationship between inflation rates and price levels in the Netherlands and the rest of the Euro-area. However, there is no reason why this model and the empirical strategy could not be applied to other Euro-area members[43]. This model and its empirical strategy can be directly extended to any 'small' member of the Euro-area. In this research we consider a country as'small' as long as its weight in the ECB's basket for measuring inflation smaller than approximately 5%. Such distinction is of course arbitrary but this can be always tested in-sample by Granger causality test.

Moreover, the empirical strategy of this paper could also be extended to larger members of the Euro-area (or any other monetary union). This could be done by running VCM and/or VECM, which would explicitly take into account the simultaneous determination of the price levels between a given country and the rest of the Euro-area. In principle even dynamic stochastic general equilibrium model could be built based on the Ca'Zorzi et al. (2005) model and empirical strategy pursued in this paper.

Forth, many scholars interested in estimating the parameters that determine inflation in the economy also tend to use crude oil prices in their estimations. However, recently some researcher pointed out that this relationship is getting weaker

[43]Or a small member of any other currency union.

(see Blanchard and Gali, 2007, Chen, 2009, for further discussion). The crude oil prices were considered to be included as a covariate in this model too, but they were found either insignificant or economically meaningless once inflation in the rest of the Euro-area was included (See table 33 below)[44]. This could be due to the fact that the Netherlands is quite progressive when it comes to alternative energy sources. More importantly this raises a question of how often could crude oil prices be included incorrectly in the estimation. It might well be that in estimation of inflation for many countries which are members of a currency union, crude oil prices are not important at all, and they just happen to be correlated with price levels in other parts of the currency union.

Table 33: Estimations of the impact of oil on the inflation rate in the Netherlands and the rest of Euro-area.

Overview of ALDR Estimations (Year-on-Year Rate). Sample: 02/2002-09/2013.						
	Year-on-Year Models			Annualized Rate Models		
Models:	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	π_{NL}	π_{NL}	π_{NL}	π_{NL}	π_{NL}	π_{NL}
π_{EA}		0.316155**	0.505560***		0.336791***	0.363664***
		[0.033, 0.600]	[0.223, 0.788]		[0.171, 0.514]	[0.15, 0.575]
		(2.201033)	(3.533208)		(3.806164)	(3.401826)
Δ Crude Oil	0.002254**	0.000515	-0.001195	0.009338***	0.006280***	0.007189***
year-on-year	[0.001, 0.004]	[-0.001, 0.002]	[-0.005, 0.003]	[0.006, 0.013]	[0.003, 0.009]	[0.004, 0.010]
	(2.537861)	(0.555584)	(-0.646779)	(5.745572)	(3.984098)	(4.780357)
Additional controls:	NO	NO	YES	NO	NO	YES
Adj. R^2	0.928	0.933	0.954	0.493	0.525	0.542
Wald F-stat.	114.627	531.407	491.925	30.607	24.968	17.360
Errors:	HAC	HAC	HAC	HAC	HAC	HAC

All estimations included constant, VAT hikes and euro dummies, one lag of π_{NL} and, in cases where π_{EA} was included, one lag of π_{EA}.
Additional controls include labour productivity and labour supply.
HAC standard errors & covariances were estimated using Bartlett kernel and Newey-West automatic lags and bandwidth.
t-statistics in (). Lower (L) and Upper (U) 95% confidence intervals are included in [L, U].
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Fifth, there is still a considerable room for improvement for the forecasts presented in the previous section. For example, better auxiliary forecasts for the independent variables could be created. Indeed, many institutions probably already have such models and thus this could be done at no additional cost. Moreover, main

[44] To be more specific, the crude oil year-on-year rate becomes insignificant once we account for inflation in the rest of the Euro-area. In the case of the annualized rate of crude oil, coefficients are significant but economically meaningless. This is so because the mean estimates of approximately 0.006 and 0.007 (or 0.6% and 0.7%) are too small to have a meaningful impact on inflation. An annualized rate of 0.7% represents only approximately 0.058% monthly increase in an inflation rate sustained during the whole year. This is negligible compared to the inflation in the rest of the Euro-area.

forecasting models itself could be improved. This work was focused on making the models as comparable as possible (even thought this was not always an option). As a result of this sometimes slightly better specifications were discarded. Furthermore, there is always room for incorporating qualitative knowledge and local know-how.

7 Conclusion

The aim of this paper has been to examine the relationship between the inflation and price levels in the Netherlands and the rest of the Euro-area. This work contributes to the current literature on optimal currency areas in several ways.

First, this paper reviews recent theoretical and empirical literature on the optimal policy in the currency areas. Special attention was put to the European Monetary Union.

Second, this paper presented a slightly modified version of the Ca'Zorzi et al. (2005) general equilibrium model of the relationship between price levels and inflation amongst two members of a monetary union. The model shows that the differences between price levels and inflation rates, in the currency area, are broadly determined by Balassa–Samuelson effects. Despite the fact that this relationship was already 'lurking' in most of the standard international macroeconomic models, this paper is one of the first one to bring attention to relationship between price levels in areas sharing the same currency and its advantages.

Third, this paper develops an empirical strategy to test predictions of the aforementioned model. This is done by estimating an ARDL model and restricted and unrestricted error correction model as well. The restricted error correction model was estimated in two steps using Johansen (2002) approach. Unrestricted error correction model was estimated using single equation approach of Pesaran et al. (2001).

These estimations provide several interesting findings. For example, in the short run, 1% increase in the EMU's inflation rate increases the Dutch inflation rate by approximately 0.5%. In the long run, 1% increase in the price level in the EMU leads to 1.1% increase in the price level in the Netherlands unconditional on the Dutch VAT hikes of 2001 and 2012, and by 0.8% conditionally on these tax hikes.

These findings are robust to changes in the model specification, sample period, empirical strategy and even to the methodology used to measure inflation (year-on-year vs. annualized monthly rate). However, this research also notices few subtle differences between using the year-on-year and annualized monthly rate. Specifically, year-on-year models are shown to have considerably better fit than annualized monthly rate. Nevertheless, at the same time models using annualized monthly rate suffered less from multicollinearity and had better-behaved errors. These differences can be of particular interest to practical forecasters and they are discussed briefly.

Fourth, even though this paper was applied specifically to the Netherlands, the theoretical and empirical strategy employed in this paper can be generalized to any 'small' member of a currency union. Conditions under which methodology, outlined in this paper, is applicable also to other countries, are discussed as well.

Fifth, this paper also tries to evaluate the potential forecasting power of the estimated model. Here it was found that there is a difference between the year-on-year and annualized rate inflation rate models. Models using the year-on-year models were found to have better performance than the annualized rate model.

Noticing the difference between the year-on-year and annualized rate is very important addition to wider research. Currently both of these ways of measuring the inflation rate are used across literature with various degrees of success. However, there is not much discussion on how the measurement of inflation rate itself can positively or adversely affect empirical findings and the models in general.

References

Altissimo, F., Benigno, P., and Palenzuela, D. R. (2011). Inflation differentials in a currency area: facts, explanations and policy. *Open Economies Review*, 22(2):189–233.

Angeloni, I. and Ehrmann, M. (2007). Euro area inflation differentials. *The BE Journal of Macroeconomics*, 7(1).

Arranz, M. A. and Escribano, A. (2000). Cointegration testing under structural breaks: A robust extended error correction model. *Oxford Bulletin of Economics and Statistics*, 62(1):23–52.

Balassa, B. (1964). The purchasing-power parity doctrine: a reappraisal. *Journal of political Economy*, 72(6):584–596.

Barnichon, R. and Nekarda, C. J. (2012). The ins and outs of forecasting unemployment: Using labor force flows to forecast the labor market. *Brookings Papers on Economic Activity*, 2012(2):83–131.

Basher, S. A. and Westerlund, J. (2008). Is there really a unit root in the inflation rate? more evidence from panel data models. *Applied Economics Letters*, 15(3):161–164.

Beck, G. W., Hubrich, K., and Marcellino, M. (2009). Regional inflation dynamics within and across euro area countries and a comparison with the United States. *Economic Policy*, 24(57):142–184.

Beyer, A., Haug, A. A., and Dewald, W. G. (2009). Structural breaks, cointegration and the fisher effect.

Blanchard, O. J. and Gali, J. (2007). The macroeconomic effects of oil shocks: Why are the 2000s so different from the 1970s? Technical report, National Bureau of Economic Research.

Breusch, T. S. (1978). Testing for autocorrelation in dynamic linear models. *Australian Economic Papers*, 17(31):334–355.

Carnot, N., Koen, V., and Tissot, B. (2011). Economic forecasting and policy.

Ca'Zorzi, M., De Santis, R. A., and Zampolli, F. (2005). Welfare implications of joining a common currency. *ECB Working Paper*.

Chen, S.-S. (2009). Oil price pass-through into inflation. *Energy Economics*, 31(1):126–133.

Cheney, E. W. and Light, W. A. (2009). *A course in approximation theory*, volume 101. American Mathematical Soc.

Clements, M. P. and Hendry, D. F. (2005). Evaluating a model by forecast performance. *Oxford Bulletin of Economics and Statistics*, 67(s1):931–956.

Cobb, C. W. and Douglas, P. H. (1928). A theory of production. *The American Economic Review*, 18(1):139–165.

Culver, S. E., Papell, D. H., et al. (1997). Is there a unit root in the inflation rate? evidence from sequential break and panel data models. *Journal of Applied Econometrics*, 12(4):435–444.

De Nederlandsche Bank (2002). Quarterly bulletin march 2002. Technical report, De Nederlandsche Bank.

Elbourne, A. and Grabska, K. (2016). Evidence on Dutch macroeconomic and sectoral productivity performance: Some stylised facts. Technical report, CPB Netherlands Bureau for Economic Policy Analysis.

Engle, R. F. and Granger, C. W. (1987). Co-integration and error correction: representation, estimation, and testing. *Econometrica: journal of the Econometric Society*, pages 251–276.

Estrella, A. (2005). Why does the yield curve predict output and inflation? *The Economic Journal*, 115(505):722–744.

Ferrero, A. (2009). Fiscal and monetary rules for a currency union. *Journal of International Economics*, 77(1):1–10.

Fratzscher, M., Duca, M. L., and Straub, R. (2014). *ECB unconventional monetary policy actions: Market impact, international spillovers and transmission channels*. International Monetary Fund.

Gambacorta, L., Hofmann, B., and Peersman, G. (2014). The effectiveness of unconventional monetary policy at the zero lower bound: A cross-country analysis. *Journal of Money, Credit and Banking*, 46(4):615–642.

Garcia, R. and Perron, P. (1996). An analysis of the real interest rate under regime shifts. *The Review of Economics and Statistics*, pages 111–125.

Glynn, J., Perera, N., and Verma, R. (2007). Unit root tests and structural breaks: a survey with applications. *Faculty of Commerce-Papers*, page 455.

Godfrey, L. G. (1978). Testing against general autoregressive and moving average error models when the regressors include lagged dependent variables. *Econometrica: Journal of the Econometric Society*, pages 1293–1301.

Granger, C. W. (1969). Investigating causal relations by econometric models and cross-spectral methods. *Econometrica: Journal of the Econometric Society*, pages 424–438.

Hamilton, J. D. (1994). *Time series analysis*, volume 2. Princeton university press Princeton.

Hyndman, R. J. and Koehler, A. B. (2006). Another look at measures of forecast accuracy. *International journal of forecasting*, 22(4):679–688.

Johansen, S. (1988). Statistical analysis of cointegration vectors. *Journal of economic dynamics and control*, 12(2):231–254.

Johansen, S. (1991). Estimation and hypothesis testing of cointegration vectors in gaussian vector autoregressive models. *Econometrica: Journal of the Econometric Society*, pages 1551–1580.

Johansen, S. (1992). Testing weak exogeneity and the order of cointegration in uk money demand data. *Journal of Policy modeling*, 14(3):313–334.

Johansen, S. (2002). A small sample correction for the test of cointegrating rank in the vector autoregressive model. *Econometrica*, 70(5):1929–1961.

Johansen, S. and Juselius, K. (1990). Maximum likelihood estimation and inference on cointegration—with applications to the demand for money. *Oxford Bulletin of Economics and statistics*, 52(2):169–210.

Krugman, P. R., Dominquez, K. M., and Rogoff, K. (1998). It's baaack: Japan's slump and the return of the liquidity trap. *Brookings Papers on Economic Activity*, 1998(2):137–205.

Kwiatkowski, D., Phillips, P. C., Schmidt, P., and Shin, Y. (1992). Testing the null hypothesis of stationarity against the alternative of a unit root: How sure are we that economic time series have a unit root? *Journal of econometrics*, 54(1):159–178.

Lombardo, G. (2006). Inflation targeting rules and welfare in an asymmetric currency area. *Journal of International Economics*, 68(2):424–442.

MacKinnon, J. G. et al. (1996). Numerical distribution functions for unit root and cointegration tests. *Journal of applied econometrics*, 11(6):601–618.

Maddala, G. S. and Kim, I.-M. (1998). *Unit roots, cointegration, and structural change*. Number 4. Cambridge University Press.

McKenzie, L. (1957). Demand theory without a utility index. *The Review of Economic Studies*, 24(3):185–189.

McQuarrie, A. D. and Tsai, C.-L. (1998). *Regression and time series model selection*. World Scientific.

Mellens, M., EUR, H. V., EUR, J. D., et al. (2014). Robust estimation of the vat pass-through in the netherlands. Technical report, CPB Netherlands Bureau for Economic Policy Analysis.

Öller, L.-E. and Barot, B. (2000). The accuracy of european growth and inflation forecasts. *International Journal of Forecasting*, 16(3):293–315.

Organisation for Economic Co-operation and Development (2001). *Measuring Productivity: Measurement of Aggregate and Industry-level Productivity Growth: OECD Manual*. Organisation for Economic Co-operation and Development.

Perron, P. (1989). The great crash, the oil price shock, and the unit root hypothesis. *Econometrica: Journal of the Econometric Society*, pages 1361–1401.

Pesaran, M. H. and Shin, Y. (1998). An autoregressive distributed-lag modelling approach to cointegration analysis. *Econometric Society Monographs*, 31:371–413.

Pesaran, M. H., Shin, Y., and Smith, R. J. (2001). Bounds testing approaches to the analysis of level relationships. *Journal of applied econometrics*, 16(3):289–326.

Phillips, P. C. and Hansen, B. E. (1990). Statistical inference in instrumental variables regression with i (1) processes. *The Review of Economic Studies*, 57(1):99–125.

Rose, A. K. (1988). Is the real interest rate stable? *The Journal of Finance*, 43(5):1095–1112.

Samuelson, P. A. (1964). Theoretical notes on trade problems. *The Review of Economics and Statistics*, pages 145–154.

Samuelson, P. A. (1974). Complementarity: An essay on the 40th anniversary of the hicks-allen revolution in demand theory. *Journal of Economic literature*, 12(4):1255–1289.

Savin, N. E. and White, K. J. (1977). The durbin-watson test for serial correlation with extreme sample sizes or many regressors. *Econometrica: Journal of the Econometric Society*, pages 1989–1996.

Stock, J. H. and Watson, M. W. (1988). Variable trends in economic time series. *The Journal of Economic Perspectives*, 2(3):147–174.

Stock, J. H. and Watson, M. W. (2007). Why has us inflation become harder to forecast? *Journal of Money, Credit and banking*, 39(s1):3–33.

Timmer, M. P., Dietzenbacher, E., Los, B., Stehrer, R., and Vries, G. J. (2015). An illustrated user guide to the world input–output database: the case of global automotive production. *Review of International Economics*, 23(3):575–605.

Varian, H. R. (1992). *Microeconomic analysis*. Norton & Company.

Verbeek, M. (2008). *A guide to modern econometrics*. John Wiley & Sons.

White, H. (1980). A heteroskedasticity-consistent covariance matrix estimator and a direct test for heteroskedasticity. *Econometrica: Journal of the Econometric Society*, pages 817–838.

Whitney K. Newey, K. D. W. (1987). A simple, positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix. *Econometrica*, 55(3):703–708.

Wu, J. C. and Xia, F. D. (2016). Measuring the macroeconomic impact of monetary policy at the zero lower bound. *Journal of Money, Credit and Banking*, 48(2-3):253–291.

Zeugner, S. (2013). Tradable vs. non-tradable: An empirical approach to the classification of sectors. Technical report, European Commision Report.

A Appendix: Testing for the Presence of Unit Root in the Inflation Series

The table 34 and 35 present the ADF and KPSS tests for the year-on-year inflation rate in the Netherlands. The reason why tests are so extensive is that they can be quite sensitive to a particular specification determined by lag selection criteria. While it is certainly true that not all criteria are equal, presenting the results based on wide range of them will give fuller picture.

Table 34: Summary of the unit root tests for year-on-year inflation rate in the Netherlands.

ADF Tests (Null: Unit Root)							
Sample	Cons.	L. Trend	No. of Lags	1% Lvl. t-Stat.	5% Lvl. t-Stat	10% Lvl. t-Stat.	t-Stat.
Full	Yes***	Yes*	24 (AIC)	-4.003005	-3.431682	-3.139538	-3.155652*
Full	Yes**	Yes*	14 (HQC, MAIC)	-4.000511	-3.430477	-3.138828	-2.671412
Full	Yes***	Yes**	13 (SIC)	-4.000316	-3.430383	-3.138772	-3.226698*
Full	Yes**	Yes*	2 (MSIC, MHQC)	-3.998280	-3.429398	-3.138192	-2.339520
Bound	Yes**	No	17 (AIC)	-3.468295	-2.878113	-2.575684	-3.033864**
Bound	Yes*	No	14 (HQC, MAIC, MHQC)	-3.468295	-2.878113	-2.575684	-2.733575*
Bound	Yes*	No	13 (SIC)	-3.468295	-2.878113	-2.575684	-3.389809**
Bound	Yes	No	2 (MSIC)	-3.468295	-2.878113	-2.575684	-2.567063
KPSS Tests (Null: Stationarity) Bandwidth: Newey-West using Bartlett kernel							
Sample	Cons.	L. Trend	Bandwidth	1% Lvl. LM-Stat.	5% Lvl. LM-Stat	10% Lvl. LM-Stat.	LM-Stat.
Full	Yes***	Yes***	11	0.216000	0.146000	0.119000	0.088068
Residual variance (no correction): 0.000126.							
HAC corrected variance (Bartlett kernel): 0.001134.							
Bound	Yes***	Yes***	10	0.216000	0.146000	0.119000	0.118406
Residual variance (no correction): 0.000100.							
HAC corrected variance (Bartlett kernel): 0.000764.							
Bandwidth: Andrews using Quadratic Spectral kernel							
Sample	Cons.	L. Trend	Bandwidth	1% Lvl. LM-Stat.	5% Lvl. LM-Stat	10% Lvl. LM-Stat.	LM-Stat.
Full	Yes***	Yes***	31.5	0.216000	0.146000	0.119000	0.052819
Residual variance (no correction): 0.000126.							
HAC corrected variance (Quadratic Spectral kernel): 0.001891.							
bound	Yes***	Yes***	20.4	0.216000	0.146000	0.119000	0.070209
Residual variance (no correction): 0.000100.							
HAC corrected variance (Quadratic Spectral kernel): 0.001289.							

Bound sample corresponds to the period 2002M02 onwards.
Linear trend was excluded if it was not significant, or if it did not seem appropriate to include it, based on a further examination of the data.
ADF confidence levels were calculated using the MacKinnon et al. (1996); one-sided p-values.
The Asymptotic critical values for the LM-stat come from the table 1 of Kwiatkowski et al. (1992).
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 35: Summary of the unit root tests for the monthly annualized inflation rate in the Netherlands

ADF Tests (Null: Unit Root)							
Sample	Cons.	L. Trend	No. of Lags	1% Lvl. t-Stat.	5% Lvl. t-Stat	10% Lvl. t-Stat.	t-Stat.
Full	Yes**	Yes*	12 (AIC, SIC, HQ)	-3.998104	-3.429313	-3.138142	-2.513944
Full	Yes	No	13 (MAIC, MSIC, MHQ)	-3.458594	-2.873863	-2.573413	-1.820818
Full, SA	Yes**	Yes*	15 (AIC, HQ)	-3.998635	-3.429570	-3.138293	-2.791682
Full, SA	Yes***	Yes*	12 (MAIC, MSIC, MHQ)	-3.998104	-3.429313	-3.138142	-2.276266
Full, SA	Yes***	Yes**	1 (SIC)	-3.996271	-3.428426	-3.137619	-12.69473***
Bound	Yes*	No	12 (AIC, SIC HQ)	-3.468295	-2.878113	-2.575684	-2.928009**
Bound, SA	Yes*	No	12 (AIC, SIC, HQ)	-3.468295	-2.878113	-2.575684	-2.869651*
Bound, SA	Yes*	No	13 (MAIC)	-3.468295	-2.878113	-2.575684	-2.848393*

KPSS Tests (Null: Stationarity) Bandwidth: Newey-West using Bartlett kernel							
Sample	Cons.	L. Trend	Bandwidth	1% Lvl. LM-Stat.	5% Lvl. LM-Stat	10% Lvl. LM-Stat.	LM-Stat.
Full	Yes***	No	46	0.739000	0.463000	0.347000	0.313553
Residual variance (no correction): 0.000126.							
HAC corrected variance (Bartlett kernel): 0.001134.							
Full, SA	Yes***	Yes*	6.05	0.216000	0.146000	0.119000	0.092892
Residual variance (no correction): 0.138562.							
HAC corrected variance (Bartlett kernel): 0.100666.							
Bound	Yes***	No	40	0.739000	0.463000	0.347000	0.217339
Residual variance (no correction): 0.344227.							
HAC corrected variance (Bartlett kernel): 0.185440.							
Bound, SA	Yes***	No	13	0.739000	0.463000	0.347000	0.220748
Residual variance (no correction): 0.169508.							
HAC corrected variance (Bartlett kernel): 0.096034.							

Bandwidth: Andrews using Quadratic Spectral kernel							
Sample	Cons.	L. Trend	Bandwidth	1% Lvl. LM-Stat.	5% Lvl. LM-Stat	10% Lvl. LM-Stat.	LM-Stat.
Full	Yes***	No	3.81	0.739000	0.463000	0.347000	0.275901
Residual variance (no correction): 0.344227.							
HAC corrected variance (Quadratic Spectral kernel): 0.210747.							
Full, SA	Yes***	Yes*	0.695	0.216000	0.146000	0.119000	0.067442
Residual variance (no correction): 0.138562.							
HAC corrected variance (Quadratic Spectral kernel): 0.138654.							
Bound	Yes***	No	3.35	0.739000	0.463000	0.347000	0.074720
Residual variance (no correction): 0.368380.							
HAC corrected variance (Quadratic Spectral kernel): 0.306982.							
Bound, SA	Yes***	No	0.873	0.739000	0.463000	0.347000	0.127015
Residual variance (no correction): 0.169508.							
HAC corrected variance (Bartlett kernel): 0.166905.							

Bound sample corresponds to the period 2002M02 onwards.
SA indicates seasonal adjustment of the series using TRAMO/SEATS.
Linear trend was excluded if it was not significant, or if it did not seem appropriate to include it, based on a further examination of the data.
ADF confidence levels were calculated using the MacKinnon et al. (1996); one-sided p-values.
The Asymptotic critical values for the LM-stat come from the table 1 of Kwiatkowski et al. (1992).
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Criteria which selected unreasonably short lag size (i.e. 0 were excluded) .

The table 36 and 37 present the ADF and KPSS tests for the year-on-year inflation rate in the Euro-area (excluding the Netherlands). As in the previous case, several estimations are reported.

Table 36: Summary of the unit root tests for year-on-year inflation rate in the EA

ADF Tests (Null: Unit Root)							
Sample	Cons.	L. Trend	No. of Lags	1% Lvl. t-Stat.	5% Lvl. t-Stat	10% Lvl. t-Stat.	t-Stat.
Full	Yes	No	12 (AIC, SIC, HQ, MAIC, MHQ)	-3.460313	-2.874617	-2.573817	-2.385415
Ext.	Yes***	Yes***	36 (AIC, MAIC)	-3.971255	-3.416382	-3.130503	-3.484830**
Ext.	Yes***	Yes***	25 (SIC, HQ, MSIC, MHQ)	-3.971049	-3.416168	-3.130376	-3.467618**
KPSS Tests (Null: Stationarity) Bandwidth: Newey-West using Bartlett kernel							
Sample	Cons.	L. Trend	Bandwidth	1% Lvl. LM-Stat.	5% Lvl. LM-Stat	10% Lvl. LM-Stat.	LM-Stat.
Full	Yes***	Yes***	11	0.216000	0.146000	0.119000	0.226593***
Residual variance (no correction): 0.804656.							
HAC corrected variance (Bartlett kernel): 7.358165.							
Ext.	Yes***	Yes***	21	0.216000	0.146000	0.119000	0.212712**
Residual variance (no correction): 0.000296.							
HAC corrected variance (Bartlett kernel): 0.005240.							
Bandwidth: Andrews using Quadratic Spectral kernel							
Sample	Cons.	L. Trend	Bandwidth	1% Lvl. LM-Stat.	5% Lvl. LM-Stat	10% Lvl. LM-Stat.	LM-Stat.
Full	Yes***	Yes***	76.2	0.216000	0.146000	0.119000	0.140409*
Residual variance (no correction): $8.05 \cdot 10^{-5}$							
HAC corrected variance (Bartlett kernel): 0.001187							
Ext.	Yes***	Yes***	139	0.216000	0.146000	0.119000	0.098864
Residual variance (no correction): 0.000296.							
HAC corrected variance (Quadratic Spectral kernel): 0.011275.							

Extended series was created using the German Mark as an precursor to the euro. The year 1996 was interpolated using Catmull-Rom spline.
Linear trend was excluded if it was not significant, or if it did not seem appropriate to include it, based on a further examination of the data.
ADF confidence levels were calculated using the MacKinnon et al. (1996); one-sided p-values.
The Asymptotic critical values for the LM-stat come from the table 1 of Kwiatkowski et al. (1992).
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 37: Summary of the unit root tests for the monthly annualized inflation in the EA (Excluding the NL).

ADF Tests (Null: Unit Root)							
Sample	Cons.	L. Trend	No. of Lags	1% Lvl. t-Stat.	5% Lvl. t-Stat	10% Lvl. t-Stat.	t-Stat.
Full	Yes	No	23 (AIC, HQ)	-3.460313	-2.874617	-2.573817	-1.561375
Full	Yes	No	12 (SIC)	-3.458845	-2.873974	-2.573472	-2.339477
Full	Yes	No	11 (MAIC, MSIC, MHQ)	-3.458719	-2.873918	-2.573443	-1.896588
Full, SA	Yes***	No	11 (AIC)	-3.458347	-2.873755	-2.573355	-3.340081**
Full, SA	Yes***	No	1 (SIC, HQ) MHQ)	-3.457173	-2.873240	-2.573080	-7.957643***
Full, SA	Yes***	Yes**	1 (SIC)	-3.996271	-3.428426	-3.137619	-12.69473***
Full, SA	Yes	No	20 (MAIC)	-3.459494	-2.874258	-2.573625	-1.736029
Full, SA	Yes**	No	10 (MSIC, MHQ)	-3.458225	-2.873701	-2.573327	-2.460097
Ext.	Yes***	No	36 (AIC)	-3.439504	-2.865470	-2.568919	-2.949027**
Ext.	Yes***	No	11 (SIC, HQ)	-3.439180	-2.865327	-2.568843	-2.956752**

KPSS Tests (Null: Stationarity) Bandwidth: Newey-West using Bartlett kernel							
Sample	Cons.	L. Trend	Bandwidth	1% Lvl. LM-Stat.	5% Lvl. LM-Stat	10% Lvl. LM-Stat.	LM-Stat.
Full	Yes***	No	145	0.739000	0.463000	0.347000	0.275969
Residual variance (no correction): 0.160707.							
HAC corrected variance (Bartlett kernel): 0.07448.							
Full, Sa	Yes***	No*	8	0.739000	0.463000	0.347000	0.463000*
Residual variance (no correction): $2.35 \cdot 10^{-06}$.							
HAC corrected variance (Bartlett kernel): $5.97 \cdot 10^{-06}$.							
Ext.	Yes***	No	19	0.739000	0.463000	0.347000	0.320864
Residual variance (no correction): 0.065317.							
HAC corrected variance (Bartlett kernel): 0.121284.							
Ext. SA	Yes***	No	8	0.739000	0.463000	0.347000	0.403643*
Residual variance (no correction): $2.35 \cdot 10^{-06}$.							
HAC corrected variance (Bartlett kernel): $5.97 \cdot 10^{-06}$.							

Bandwidth: Andrews using Quadratic Spectral kernel							
Sample	Cons.	L. Trend	Bandwidth	1% Lvl. LM-Stat.	5% Lvl. LM-Stat	10% Lvl. LM-Stat.	LM-Stat.
Full	Yes***	No	1.01	0.739000	0.463000	0.347000	0.126132
Residual variance (no correction): 0.160707.							
HAC corrected variance (Quadratic Spectral kernel): 0.162973.							
Full, SA	Yes***	No*	6.86	0.739000	0.463000	0.347000	0.403657*
Residual variance (no correction): $2.35 \cdot 10^{-06}$.							
HAC corrected variance (Quadratic Spectral kernel): $5.97 \cdot 10^{-06}$.							
Ext.,	Yes***	No*	9.72	0.739000	0.463000	0.347000	0.463816**
Residual variance (no correction): 0.065317.							
HAC corrected variance (Quadratic Spectral kernel): 0.083903.							
Ext., SA	Yes***	No	9.4	0.739000	0.463000	0.347000	0.462277*
Residual variance (no correction): 0.019511.							
HAC corrected variance (Quadratic Spectral kernel): 0.081580.							

Extended series was created using the German Mark as an precursor to the euro. The year 1996 was interpolated using Catmull-Rom spline.
SA indicates seasonal adjustment of the series using TRAMO/SEATS.
Linear trend was excluded if it was not significant, or if it did not seem appropriate to include it, based on a further examination of the data.
ADF confidence levels were calculated using the MacKinnon et al. (1996); one-sided p-values.
The Asymptotic critical values for the LM-stat come from the table 1 of Kwiatkowski et al. (1992).
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

B Appendix: Unit root Tests of Control Variables

The table 38 reports ADF unit root tests for the year-on-year and monthly annualized rates of other control variables. Here the tests are less extensive because stationarity of these variables is not as controversial as in the case of inflation. All variables are seasonally and calendar adjusted.

Some of the variables are significant only at 10% confidence interval, but this is almost certainly due to lack of power on the side of ADF test. Table 39 shows that we cannot reject the null at the 5% level for the same variables as in the ADF case.

Table 38: ADF unit root tests for control variables.

ADF Tests (Null: Unit Root)							
Variable	Cons.	L. Trend	No. of Lags	1% Lvl. t-Stat.	5% Lvl. t-Stat	10% Lvl. t-Stat.	t-Stat.
Δda_{NL} (yoy)	Yes	No	12 (AIC)	-3.480425	-2.883408	-2.578510	-2.710540**
Δda_{EA} (yoy)	Yes**	No	20 (AIC)	-4.000511	-3.430477	-3.138828	-2.671412*
Δl_{NL} (yoy)	Yes	No	24 (AIC)	-3.468072	-2.878015	-2.575632	-2.588367*
Δl_{EA} (yoy)	Yes**	Yes**	11 (AIC)	-4.024452	-3.442006	-3.145608	-3.403966*
Δda_{NL} (ma)	Yes*	No	20 (AIC)	-3.465014	-2.876677	-2.574917	-3.413841**
Δda_{EA} (ma)	Yes	No	8 (AIC)	-3.472813	-2.880088	-2.576739	-3.192144**
Δl_{NL} (ma)	Yes	No	10 (AIC)	-3.463235	-2.875898	-2.574501	-2.840326*
Δl_{EA} (ma)	Yes	No	2 (AIC)	-3.473382	-2.880336	-2.576871	-3.019696**

"yoy" denotes year-on-year rate and "ma" monthly annualized rate
Linear trend was excluded if it was not significant, since it is not apriori clear wether the series has one.
ADF confidence levels were calculated using the MacKinnon et al. (1996); one-sided p-values.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 39: KPSS unit root test of selected control variables.

KPSS Tests (Null: Stationary)							
Variable	Cons.	L. Trend	Bandwidth	1% Lvl. LM-Stat.	5% Lvl. LM-Stat	10% Lvl. LM-Stat.	LM-Stat.
Δda_{EA} (yoy)	Yes***	No	10	0.739000	0.463000	0.347000	0.108666
Δl_{NL} (yoy)	Yes***	Yes***	10	0.216000	0.146000	0.119000	0.134380*
Δl_{EA} (yoy)	Yes***	Yes***	9	0.216000	0.146000	0.119000	0.075102
Δl_{NL} (ma)	Yes***	Yes***	4	0.216000	0.146000	0.119000	0.135798*

"yoy" denotes year-on-year rate and "ma" monthly annualized rate
Linear trend was excluded if it was not significant, since it is not apriori clear wether the series has one.
The tests were performed using Bartlett kernel and Newey-West automatic bandwidth .
The Asymptotic critical values for the LM-stat come from the table 1 of Kwiatkowski et al. (1992).
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

C Appendix: Structural Factors by Sectoral Breakdown

This appendix provides a detailed description of the NACE industry classification. Table 40 follows the NACE Rev. 2 Statistical classification of economic activities in the European Community.

Table 40: Explanatory Notes to the NACE codes

NACE	Explanation
TOTAL	Total - All NACE activities.
A	Agriculture, forestry and fishing.
B-E	Industry (except construction).
C	Manufacturing.
F	Construction.
G-I	Wholesale and retail trade, transport, accomodation and food service activities. accomodation and food service activities.
J	Information and communication.
K	Financial and insurance activities.
L	Real estate activities.
M_N	Professional, scientific and technical activities; administrative and support service activities.
O-Q	Public administration, defence, education, human health and social work activities.
R-U	Arts, entertainment and recreation; other service activities; activities of household and extra-territorial organizations and bodies.

D Appendix: Hours Worked

Table 41: Descriptive statistics of hours worked by sector (in thousands) in the Netherlands.

	A	B-E	C	F	G-I	J	K	L	M_N	O-Q	R-U
Mean	116860.3	389930.4	360240.2	200650.1	737437.5	98137.27	103341.6	25998.45	434686.0	647192.2	139411.5
Median	116966.0	381885.0	352989.0	201483.0	740533.0	102702.0	104359.0	25943.00	429689.0	662330.0	136529.0
Max.	135991.0	425238.0	394598.0	216153.0	764147.0	112437.0	112587.0	29390.00	506798.0	730391.0	157120.0
Min.	102582.0	354076.0	323702.0	143341.0	678674.0	60263.00	90867.00	22177.00	315524.0	533043.0	117663.0
S. D.	8267.044	25630.33	25558.62	11456.55	18067.47	13052.42	5398.250	1405.294	44119.89	60550.90	10208.49
Skew.	0.402568	0.045765	0.015494	−2.617626	−0.967737	−1.465771	−0.560585	0.135083	−0.533770	−0.214308	0.143132
Kurt.	2.438574	1.350613	1.380290	13.40239	3.646121	4.184660	2.393621	3.119539	2.843154	1.716017	2.271830

Table 42: Descriptive statistics of hours worked by sector (in thousands) in the European Monetary Union (excluding the NL).

	A	B-E	C	F	G-I	J	K	L	M_N	O-Q	R-U
Mean	2659811.	9235675.	8547211.	4399626.	13878369	1425224.	1476665.	434187.6	5812588.	9642946.	3437013.
Median	2652854.	9357181.	8671694.	4375839.	13931082	1438847.	1470432.	449998.0	5907040.	10754752	3515123.
Max.	3093345.	10033614	9352280.	4935913.	14470444	1523812.	1517399.	477418.0	6680524.	11604129	3673151.
Min.	2342688.	8458671.	7748339.	3721728.	12529967	1245138.	1405806.	383723.0	4528790.	0.000000	2986525.
S. D.	198521.5	442219.2	461687.5	309889.1	501509.4	72810.51	30513.66	31133.61	661007.9	3504353.	211839.4
Skew.	0.302830	−0.176376	−0.217339	−0.246446	−1.134911	−0.58755	−0.48791	−0.30714	−0.27399	−2.35036	−0.68305
Kurt.	2.351044	1.771749	1.720775	2.848331	4.010484	2.548802	2.651819	1.483438	1.683024	6.716189	2.124696

(a) Netherlands **(b)** EA (Excluding NL)

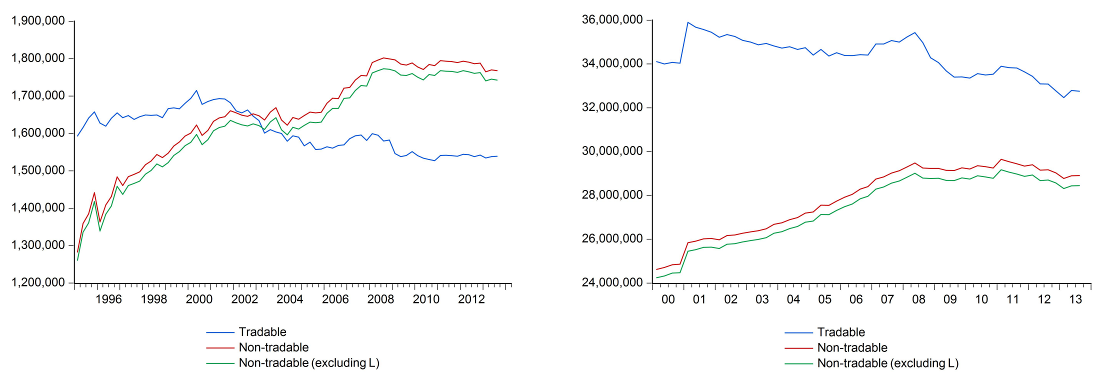


Figure 20: Plots of hours worked in tradable and non-tradable sector in the Netherlands and EA (excluding the NL).

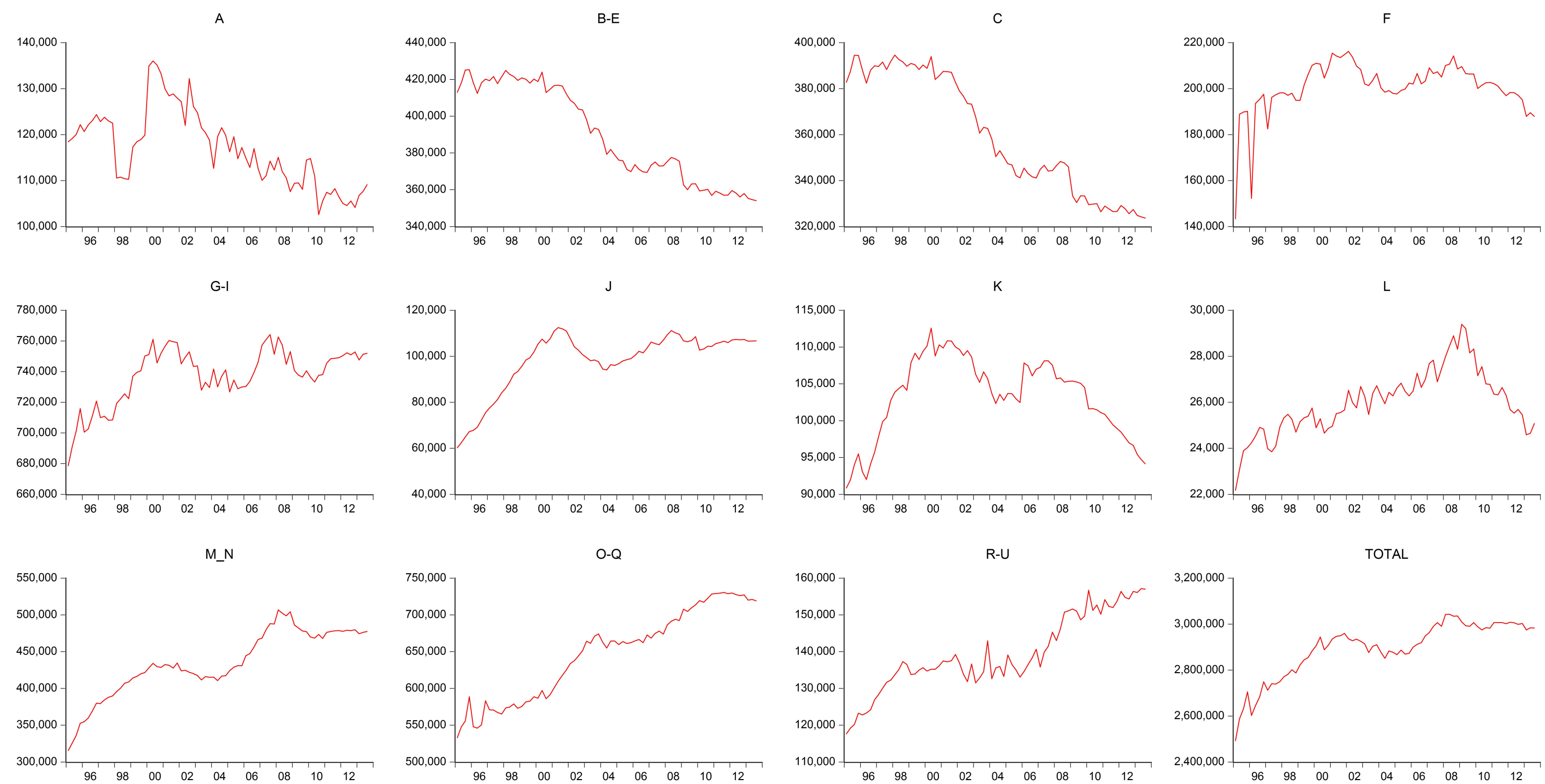


Figure 21: Labour supply in the Netherlands, measured by hours worked (in thousands) by sector (see table 40 for explanations of NACE codes.)

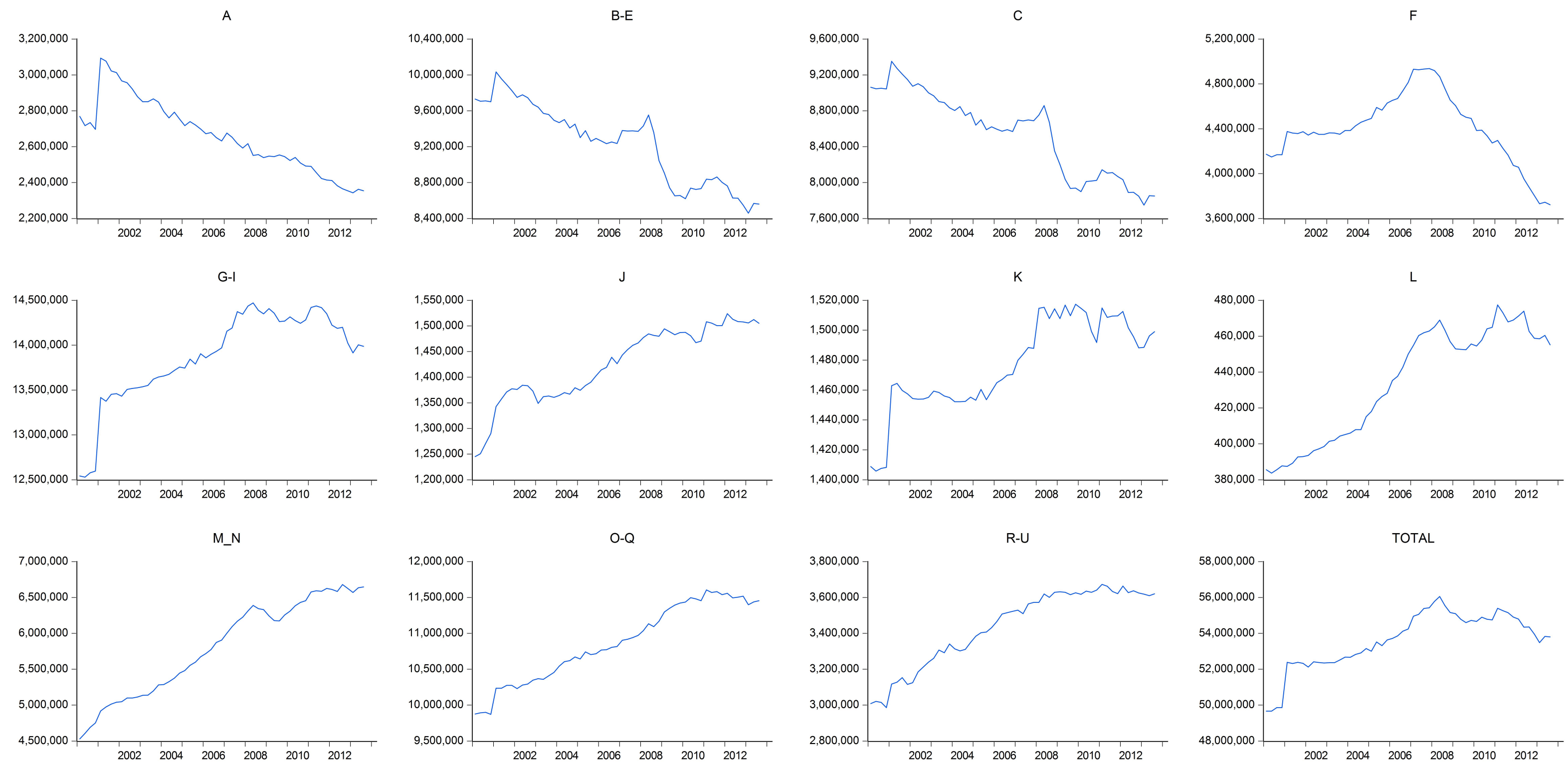


Figure 22: Labour supply in the European Monetary Union (excluding the Netherlands and with changing composition), measured by hours worked (in thousands) by sector (see table 40 for explanations of NACE codes.)

E Appendix: Real Gross Value Added

Table 43: Descriptive statistics of real gross value added by sector (in millions) in the Netherlands.

	A	B-E	C	F	G-I	J	K	L	M_N	O-Q	R-U
Mean	2501.321	22680.41	16228.97	7745.306	25876.39	5933.155	10012.70	7719.786	18127.89	27887.62	3443.604
Median	2488.959	23080.67	16555.32	7690.325	26337.90	6393.992	10288.92	7523.483	18227.12	27772.55	3496.705
Max.	2850.620	25938.39	18379.10	9108.121	30657.35	7896.751	12041.70	9349.062	21980.75	31761.11	3765.042
Min.	2082.471	19412.34	13061.58	6256.051	18465.62	2404.699	6802.718	6755.712	13481.52	22779.26	2800.633
S. D.	188.6778	1523.693	1374.381	602.6792	3194.645	1560.412	1455.525	664.2658	1986.436	3060.767	255.0246
Skew.	−0.136038	−0.606690	−0.70803	0.06517	−0.547535	−0.87042	−0.43649	1.084579	−0.371064	−0.14994	−0.86061
Kurt.	2.027651	2.550947	2.637307	2.822001	2.388930	2.623884	2.116835	3.376953	2.713371	1.639740	2.781217

Table 44: Descriptive statistics of Real gross value added by sector (in millions) in the European Monetary Union (Excluding the Netherlands).

	A	B-E	C	F	G-I	J	K	L	M_N	O-Q	R-U
Mean	31023.21	372777.9	313972.8	113817.6	354743.1	77164.42	91295.95	212611.5	188214.9	359120.8	66969.72
Median	31128.68	376835.2	314247.2	114636.7	370540.1	77296.82	90370.17	215490.9	193268.9	359598.9	68530.07
Max.	34699.40	417216.3	355113.3	128787.3	398683.0	108484.5	101338.6	246955.1	216276.6	400188.3	74400.51
Min.	26931.56	306788.0	254488.5	98710.26	281695.2	39467.22	74694.80	168868.5	144281.2	310717.0	55849.68
S. D.	2024.077	31122.60	28390.49	8630.773	34150.37	20701.63	8168.795	24396.39	20966.44	27858.40	5677.439
Skew.	−0.146917	−0.58658	−0.48199	−0.20986	−0.76322	−0.23283	−0.46945	−0.32412	−0.66622	−0.21614	−0.49477
Kurt.	1.995374	2.272171	2.191754	2.017671	2.308022	1.826995	1.935019	1.738587	2.274153	1.695455	1.860933

(a) Netherlands **(b)** EA (Excluding NL)

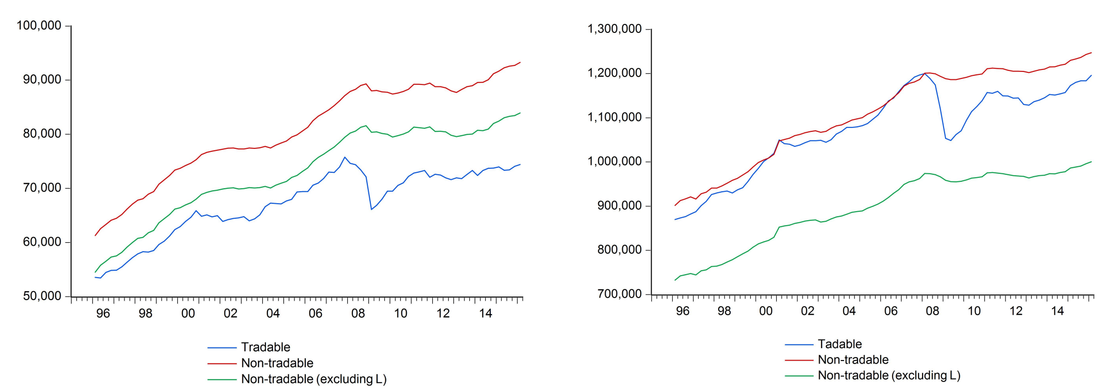


Figure 23: Plots of real gross value added in tradable and non-tradable sectors in the Netherlands and the Euro-area (excluding the Netherlands).

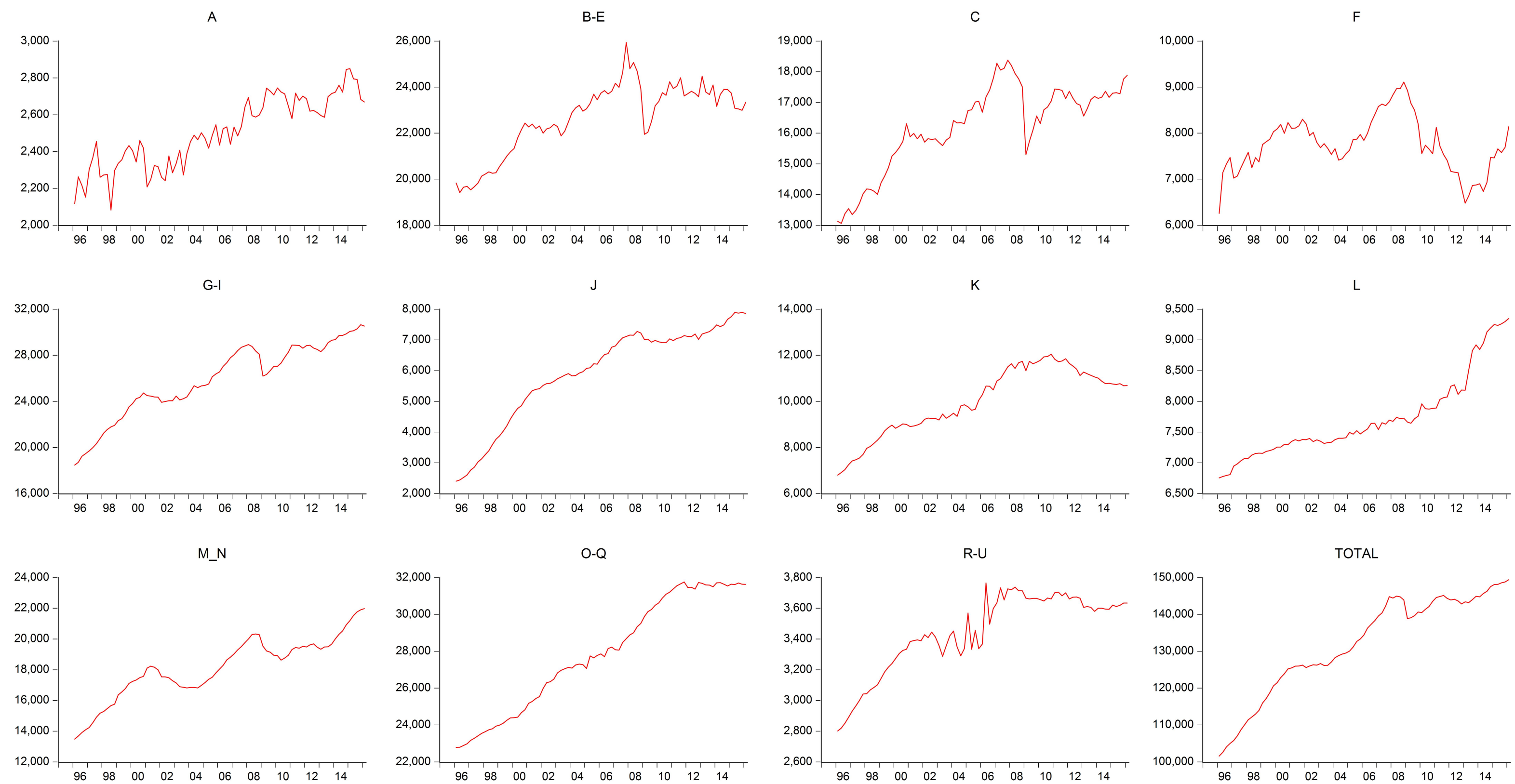


Figure 24: Real gross value added in the Netherlands measured using euros (in millions) by sector (see table 40 for explanations of NACE codes.)

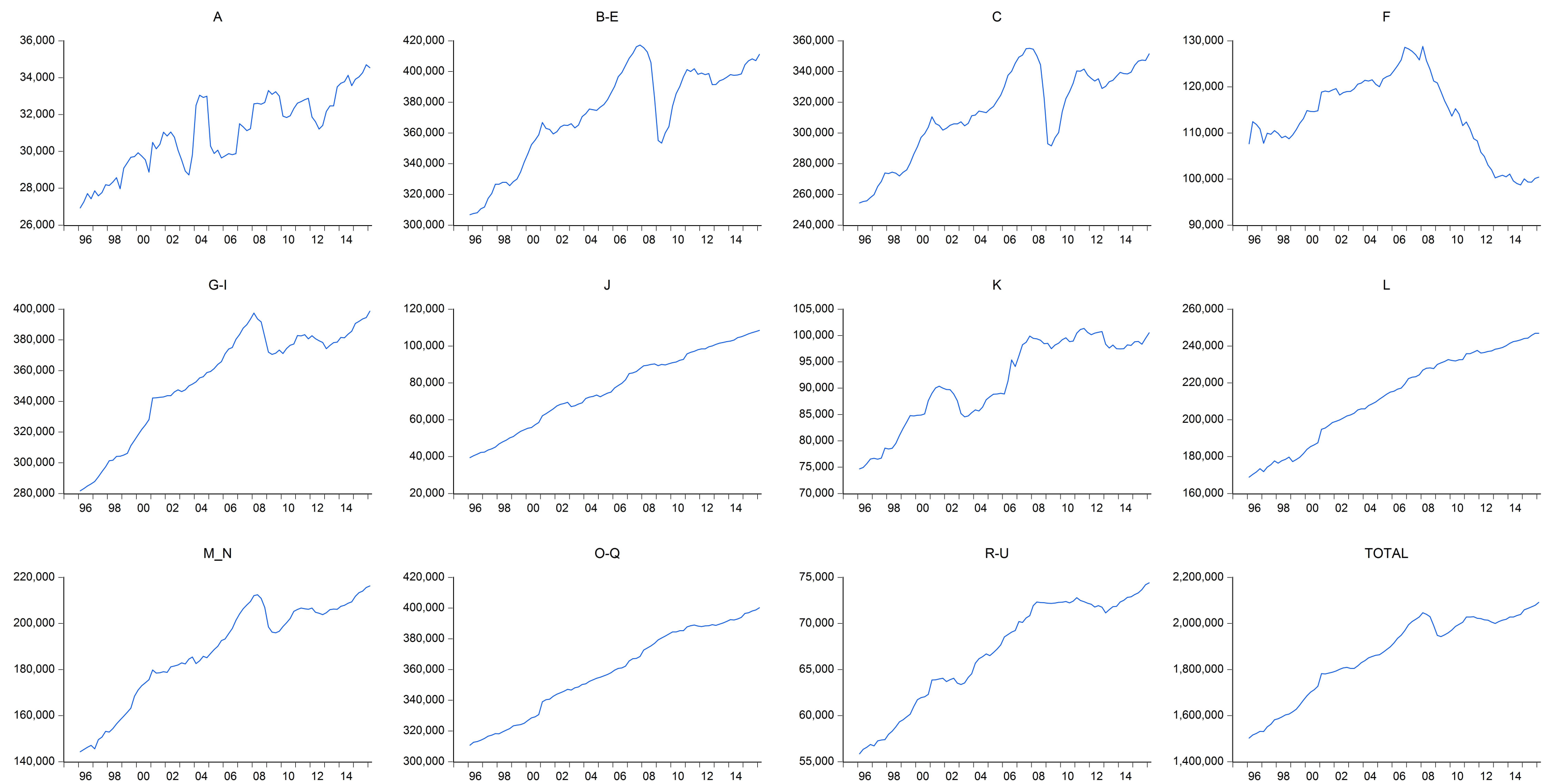


Figure 25: Real gross value added in the European Monetary Union (Excluding Netherlands) measured using euros (in millions) by sector (see table 40 for explanations of NACE codes.)

F Appendix: Labour Productivity

Table 45: Descriptive statistics of labour productivity by sector in the Netherlands.

	A	B-E	C	F	G-I	J	K	L	M_N	Q-O	R-U
Mean	21.30349	58.58065	45.23826	38.61418	34.15699	56.03699	95.49052	287.9150	40.20184	41.88652	24.34739
Median	20.87562	60.76740	46.56754	38.21329	34.73715	62.13146	93.83877	283.7195	40.51481	41.77039	24.33904
Max.	26.44434	69.54272	53.36819	43.68697	38.75104	68.14135	119.7232	352.0498	42.45564	44.00870	27.58657
Min.	16.99787	46.59136	33.82982	34.48988	26.35597	34.50819	73.10270	260.7820	37.09488	39.40905	22.46431
S. D.	2.724386	7.414363	6.524799	1.956721	3.609149	11.68365	15.79335	17.08709	1.108702	0.976979	1.120905
Skew.	0.121817	−0.293659	−0.348264	0.518264	−0.457189	−0.650736	0.213027	1.775256	−0.932250	0.221383	0.381071
Kurt.	1.682773	1.599190	1.678260	2.935250	2.076683	1.798016	1.540846	6.577467	3.595731	2.642397	2.840324

Table 46: Descriptive statistics of labour productivity in the European Monetary Union (Excluding the NL).

	A	B-E	C	F	G-I	J	K	L	M_N	O-Q	R-U
Mean	11.84844	41.50954	37.79740	26.70540	26.31705	56.45907	63.32384	501.2063	33.61372	33.51475	19.92251
Median	11.81156	42.06643	37.82053	26.68851	26.22709	57.29878	64.28608	502.9542	33.86190	33.53026	19.89025
Max.	13.79159	46.28972	42.63401	27.71748	27.53717	67.73531	67.69856	525.5163	37.77638	34.10759	20.86153
Min.	9.796619	35.59730	32.11458	25.50123	25.37430	44.46711	57.98476	477.1998	30.60234	33.09320	19.21474
S. D.	1.222980	3.406138	3.410742	0.616494	0.617552	6.909687	2.940663	10.58104	1.988739	0.227353	0.322546
Skew.	−0.017443	−0.191981	−0.102399	0.006820	0.219818	−0.110709	−0.225711	−0.306474	0.179482	0.416555	0.712124
Kurt.	1.628698	1.626382	1.543414	1.779340	1.787381	1.784312	1.607870	2.611497	1.994573	3.381196	3.715081

(a) Netherlands **(b)** EA (Excluding NL)

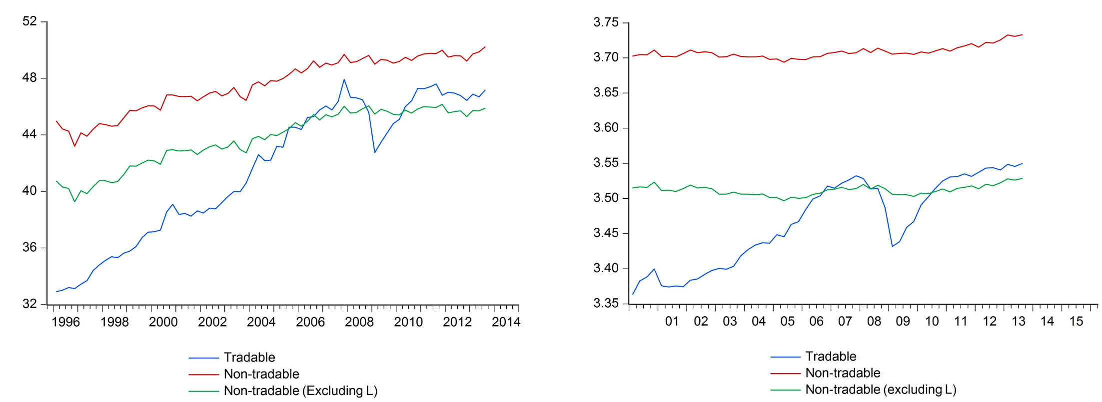


Figure 26: Plots of log of labour productivity in tradable and non-tradable sector in the Netherlands and EA (excluding the NL).

Figure 27: Log of labour productivity in the Netherlands by sector (see table 40 for explanations of NACE codes.)

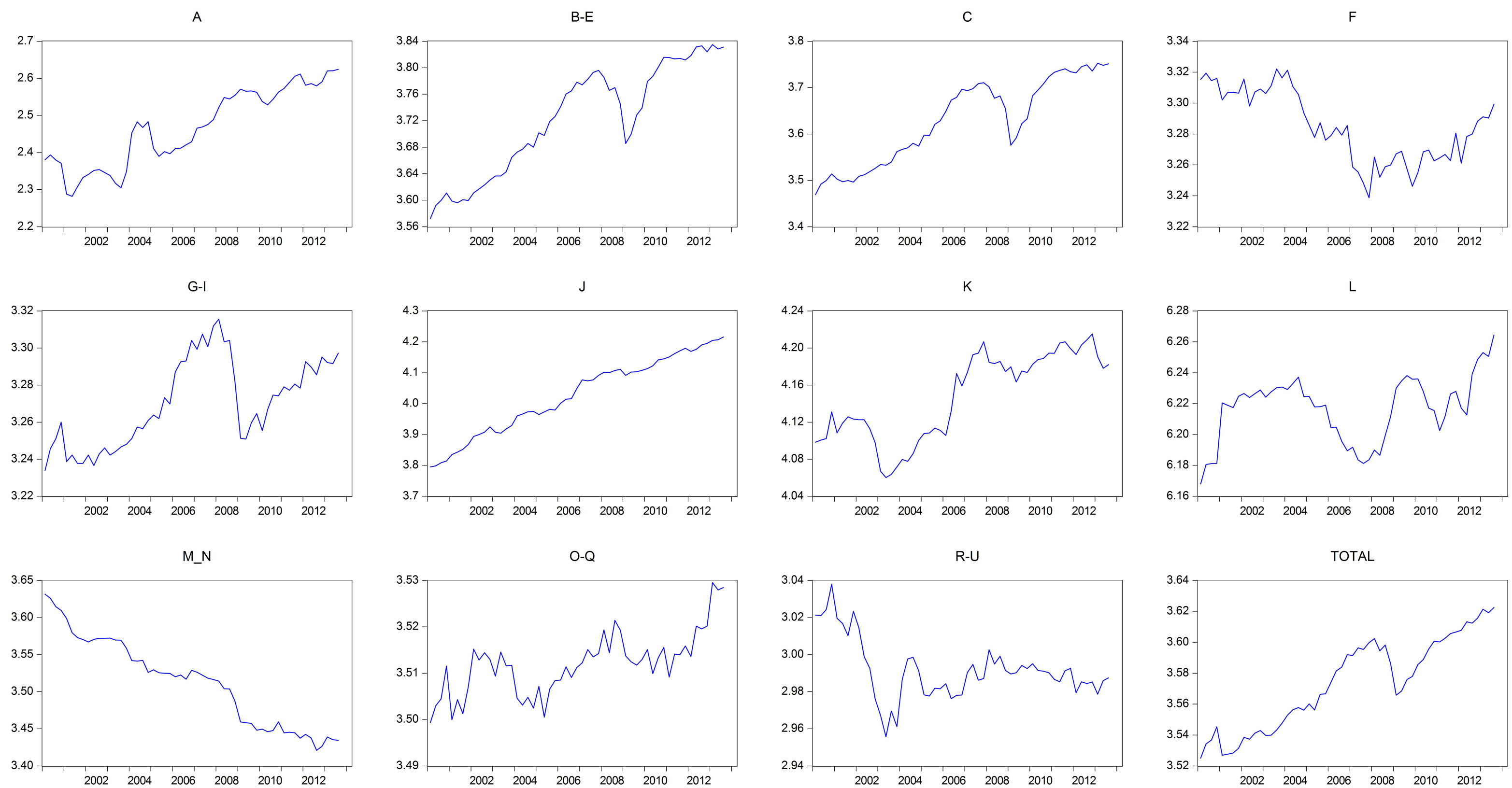


Figure 28: Log of labour productivity in the European Monetary Union (excluding the Netherlands) by sector (see table 40 for explanations of NACE codes.)

G Appendix: Further Robustness Checks

This section provides extra robustness checks to the estimations presented in section 4.1. These robustness checks take advantage of the fact that excluding structural parameters allows to extend the sample from January 2001 - September 2013 to January 1997 to April 2016 in the case of year-on-year models. In the case of annualized monthly rate models, this allows us to extend the database from January 2001 - September 2013 to February 2000 -September 2013 and to February 1996 to June 2016.

Table 47: Summary of the estimations of the relationship between the year-on-year inflation rate in the Netherlands and the rest of the Euro-area.

Overview of Furtherer Robustness Checks (Year-on-Year Rate). Sample: 01/1997-03/2016.			
Models:	(3)	(4)	(5)
Dependent Variable	π_{NL}	π_{NL}	π_{NL}
π_{EA}	0.452237***	0.551391***	0.623310***
	[0.233, 0.672]	[0.283, 0.780]	[0.159, 1.088]
	(4.056460)	(4.051062)	(2.644888)
$L^1\pi_{EA}$	-0.278990**		
	[-0.506, -0.052]		
	(-2.425257)		
$L^1\pi_{NL}$	0.705625***		
	[0.595, 0.817]		
	(12.52203)		
Dummy	0.011231***	0.033279***	
2001 VAT	[0.007, 0.016]	[0.030, 0.037]	
	(4.899250)	(19.98396)	
Dummy	0.004381***	0.019422***	
2002 Euro	[0.002, 0.007]	[0.014, 0.025]	
	(3.146763)	(6.972323)	
Dummy	0.004510***	0.014500***	
2012 VAT	[0.003, 0.007]	[0.012, 0.018]	
	(4.425030)	(9.405840)	
Constant	0.00186**	0.006854**	0.009151***
	[0.001, 0.004]	[0.001, 0.013]	[0.003, 0.015]
	(2.731771)	(2.391005)	(3.106449)
Adj. R^2	0.879607	0.739357	0.220047
SE of reg.	0.004195	0.006160	0.010655
$\sum\varepsilon^2$	0.003924	0.008574	0.025999
Durbin-Wat. stat.	1.809823	0.542480	0.187983
Wald F-stat.	512.0151	682.9584	6.995433
Errors:	HAC	HAC	HAC

The μ and σ of dep. var. during the sample period were 0.019498 and 0.012065 respectively
HAC standard errors & covariances were estimated using Bartlett kernel and Newey-West automatic lags and bandwidth.
t-statistics in (). Lower (L) and Upper (U) 95% confidence intervals are included in [L, U].
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 48: Summary of the estimations of the relationship between the annualized inflation rate in NL and EA.

Overview of Furtherer Robustness Checks (Anualized Rate)						
Models:	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	π_{NL}	π_{NL}	π_{NL}	π_{NL}	π_{NL}	π_{NL}
Sample:	02/2000 -09/2013	02/2000 -09/2013	03/1996 -06/2016	02/1996 -06/2016	02/1996 -06/2016	02/2000 -09/2013
π_{EA}	0.508585***	0.480087***	0.647259***	0.575096***	0.357208	0.502883***
	[0.323, 0.695]	[0.310, 0.650]	[0.370, 0.925]	[0.386, 0.764]	[0.017, 0.810]	[0.313, 0.692]
	(5.398394)	(5.584544)	(4.595089)	(5.997661)	(1.553556)	(5.258933)
$L^1\pi_{EA}$	-0.102764*	-0.130493***	-0.229153*			-0.099021*
	[-0.215, 0.010]	[-0.228, -0.033]	[-0.471, 0.013]			[-0.201, 0.003]
	(-1.802683)	(-2.650108)	(-1.864631)			(-1.911518)
$L^1\pi_{NL}$	0.081420**	0.101028***	-0.013293			0.077861**
	[0.007, 0.156]	[0.037, 0.166]	[-0.124, 0.100]			[0.015, 0.141]
	(2.160025)	(3.090613)	(-0.237361)			(2.447085)
Δda_{NL}	0.021084					0.023373
	[-0.036, 0.079]					[-0.031, 0.078]
	(0.725846)					(0.846825)
Δda_{EA}	-0.085778*					-0.087546*
	[-0.185, 0.013]					[-0.182, 0.007]
	(-1.714456)					(-1.824026)
Δl_{NL}	-0.006025	-0.012208				
	[-0.084, 0.072]	[-0.091, 0.067]				
	(-0.152375)	(-0.306391)				
Δl_{EA}	-0.012547	-0.010507				
	[-0.080, 0.055]	[-0.080, 0.059]				
	(-0.369261)	(-0.300301)				
Dummy M 2001 VAT	0.212188***	0.213265***	0.228731***	0.220381***		0.209270***
	[0.195, 0.300]	[0.195, 0.232]	[0.200, 0.260]	[0.202, 0.239]		[0.193, 0.226]
	(23.78678)	(22.82149)	(15.50351)	(23.24002)		(25.12468)
Dummy M 2002 Euro	0.097029***	0.096802***	0.099408***	0.098184***		0.097726***
	[0.093, 0.101]	[0.094, 0.105]	[0.094, 0.105]	[0.093, 0.104]		[0.093, 0.103]
	(36.91105)	(48.17939)	(33.51354)	(35.15308)		(38.68033)
Dummy M 2012 VAT	0.101857***	0.105319***	0.108502***	0.104614***		0.102087***
	[0.099, 0.105]	[0.101, 0.110]	[0.102, 0.115]	[0.101, 0.108]		[0.099, 0.106]
	(59.78295)	(49.70835)	(32.60864)	(54.34682)		(5.095167)
Constant	0.010794***	0.010711***	0.010329***	0.007600**	0.012832**	0.010832***
	[0.006, 0.016]	[0.006, 0.016]	[0.006, 0.015]	[0.000, 0.015]	[0.002, 0.024]	[0.006, 0.016]
	(4.054018)	(4.046277)	(4.375409)	(2.058209)	(2.278235)	(4.106531)
Adj. R^2	0.473563	0.468438	0.138623	0.137297	0.017324	0.480093
SE of reg.	0.019982	0.020079	0.041916	0.041864	0.044680	0.019858
$\sum \varepsilon^2$	0.061090	0.062491	0.416405	0.420620	0.485103	0.061121
Durbin-Wat. stat.	1.928503	1.909187	1.934574	1.968400	1.942807	1.921868
F-Stat.	15.66290	18.95544	7.517766	10.70802	5.301499	19.81468
Errors:	HAC	HAC	HAC	HAC	HAC	HAC
μ dep. var.	0.022365	0.022365	0.018557	0.018579	0.018579	0.022365
σ dep. var.	0.027540	0.027540	0.045163	0.045072	0.045072	0.027540

HAC standard errors & covariances were calculated using prewhitening with lags based on AIC (with max lag = 5), Bartlett kernel and Newey-West bandwidth.

t-statistics is included in ().

Lower (L) and Upper (U) 95% confidence intervals are included in [L, U].

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

H Appendix: Overview of Results from White Heteroskedasticity Tests

Table 49: Summary of the estimations of the relationship between the annualized inflation rate in NL and EA.

Overview of results from the White heteroskedasticity test.						
Models:	(1)	(2)	(3)	(4)	(5)	(6)
Dependent	π_{NL}	π_{NL}	π_{NL}	π_{NL}	π_{NL}	π_{NL}
F-stat.						
Obs*R-squared						
Obs*R-squared						
Scaled explained SS						
Obs*R-squared						

All tests include the White cross-term. White's (1980)
Null hypothesis: errors are homoskedastic.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

I Appendix: Plots of Residuals from the 'Short-run' Estimations

The figure 29 and 30 show plots of residuals from the first 'Short-run' models.

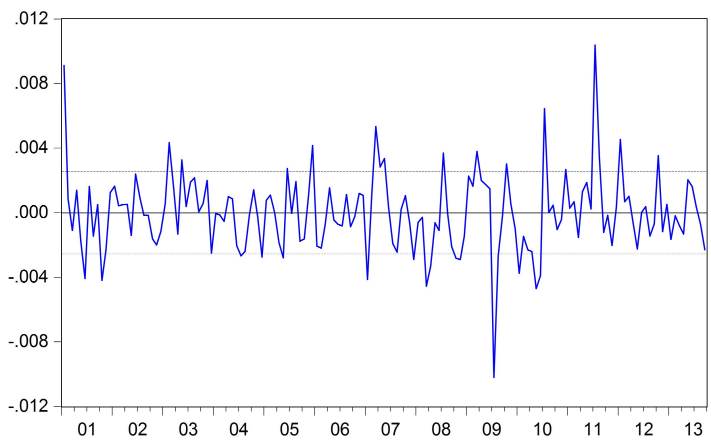


Figure 29: Residuals of year-on-year model (1).

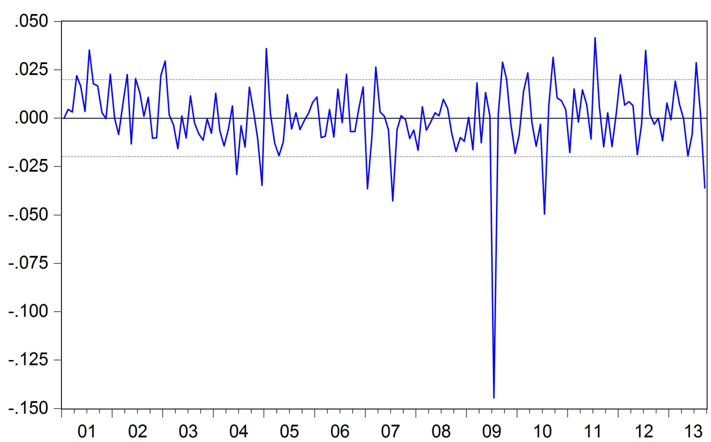


Figure 30: Residuals of annualized monthly rate model (1).

J Appendix: Vector Error Correction

The results of auxiliary vector error correction for the Johansen test.

Table 50: Vector Error Correction Estimates. Sample: 01/2002 - 06/2016

Error Correction:	$\Delta(p_{NL})$	$\Delta(p_{EA})$
CointEq1	-0.115484	-0.040246
	(0.03340)	(0.01229)
	[-3.45804]	[-3.27499]
$\Delta(L^1 p_{NL})$	0.037151	0.034257
	(0.07910)	(0.02911)
	[0.46970]	[1.17700]
$\Delta(L^2 p_{NL})$	-0.235859	0.003804
	(0.07704)	(0.02835)
	[-3.06145]	[0.13417]
$\Delta(L^3 p_{NL})$	-0.008273	0.014284
	(0.08477)	(0.03119)
	[-0.09759]	[0.45791]
$\Delta(L^1 p_{EA})$	-0.411622	0.111181
	(0.22088)	(0.08128)
	[-1.86353]	[1.36789]
$\Delta(L^2 p_{EA})$	0.376951	0.164096
	(0.22284)	(0.08200)
	[1.69156]	[2.00116]
$\Delta(L^3 p_{EA})$	-0.079760	-0.020897
	(0.22475)	(0.08270)
	[-0.35489]	[-0.25268]
C	0.001712	0.000948
	(0.00053)	(0.00020)
	[3.22288]	[4.84993]
Adj. R-squared	0.126436	0.177516
Sum sq. resids	0.002533	0.000343
S.E. equation	0.003906	0.001437
F-statistic	4.577030	6.334057
Log likelihood	722.0674	896.0215
Akaike AIC	-8.207672	-10.20714
Schwarz SC	-8.062428	-10.06190
Mean dependent	0.001302	0.001365
S.D. dependent	0.004179	0.001585
Determinant resid covariance (dof adj.)		$3.08E-11$
Determinant resid covariance		$2.81E-11$
Akaike information criterion		-18.41396
Schwarz criterion		-18.08716

Standard errors in ( ) & t-statistics in [ ]

The figure 51 plots the roots of AR characteristic polynomial and table figure 31 shows their values. VEC specification imposes 1 unit root on the model, and the table 51 shows that indeed only one of the roots lies on the unit cycle. This implies that the model is dynamically stable accounting for the cointegration.

Table 51: Roots of Characteristic Polynomial (3 lags).

Root	Modulus
1.000000	1.000000
0.961071	0.961071
0.036283 - 0.508022i	0.509317
0.036283 + 0.508022i	0.509317
0.375398	0.375398
-0.266065 - 0.060694i	0.272900
-0.266065 + 0.060694i	0.272900
0.188265	0.188265

VEC specification imposes 1 unit root(s).

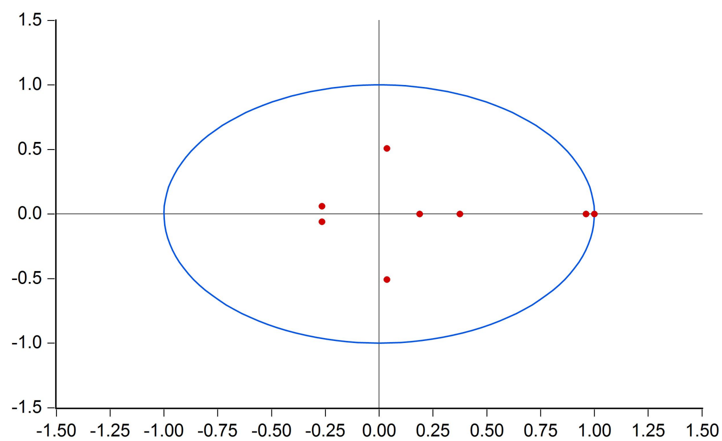


Figure 31: The Inverse Roots of AR Characteristic Polynomial

K Appendix: Bound Tests

This section presents results from the 'bounds tests' for cointegration. These tests are in essence Wald tests, testing the hypothesis that coefficients of all cointegrating variables are jointly equal to zero against the alternative that they are not. However, standard critical values do not apply anymore, thus the test statistics need to be compared to Pesaran et al. (2001) critical values instead (See their table CI). Tables 52 and 53 show results of these tests.

Table 52: Bounds Tests for year-on-year models

Wald Test: year-on-year model (1)

Test Statistic	Value	df
F-statistic	24.97238	(6, 139)
Chi-square	149.8343	6

Wald Test: year-on-year model (2)

Test Statistic	Value	df
F-statistic	2.886203	(6, 127)
Chi-square	17.31722	6

$H_0 : \mu = \varphi = \tau = \psi = \varrho = \psi = 0$

Wald Test: year-on-year model (3)

Test Statistic	Value	df
F-statistic	7.924690	(2, 129)
Chi-square	15.84938	2

$H_0 : \mu = \varphi = 0$

Table 53: Bounds tests for annualized monthly rate models.

Wald Test: annualized monthly rate model (1)

Test Statistic	Value	df
F-statistic	22.65633	(6, 139)
Chi-square	135.9380	6

Wald Test: annualized monthly rate model (2)

Test Statistic	Value	df
F-statistic	19.56028	(6, 127)
Chi-square	117.3617	6

$H_0 : \mu = \varphi = \tau = \psi = \varrho = \psi = 0$

Wald Test: year-on-year model (3)

Test Statistic	Value	df
F-statistic	4.048339	(2, 167)
Chi-square	8.096678	2

$H_0 : \mu = \varphi = 0$

L Appendix: Auxiliary Results for Forecasting Models

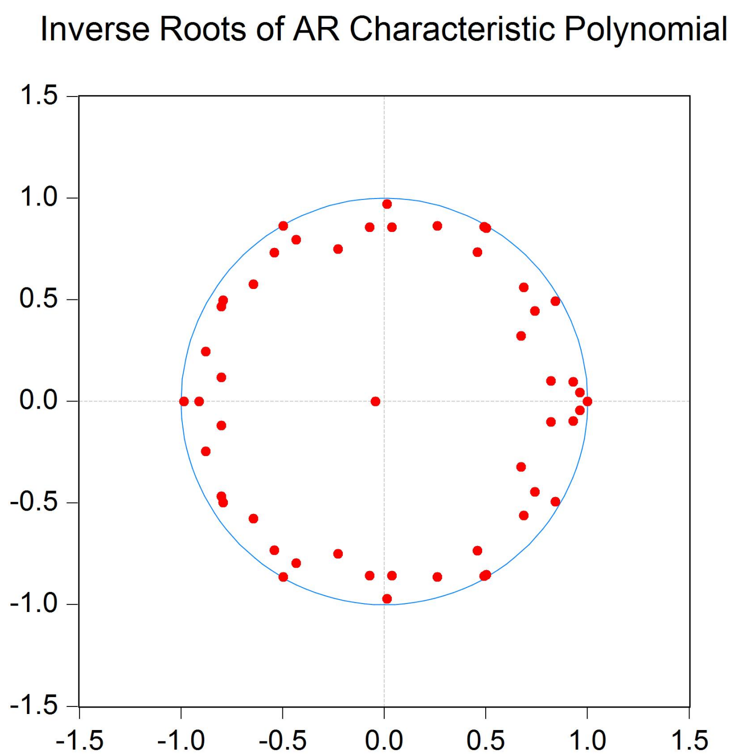


Figure 32: The Inverse Roots of AR Characteristic Polynomial

Table 54: Roots of Characteristic Polynomial (13 lags)

Root	Modulus
1.000000	1.000000
1.000000 - 1.72e-14i	1.000000
1.000000 + 1.72e-14i	1.000000
-0.497611 - 0.863987i	0.997041
-0.497611 + 0.863987i	0.997041
0.502688 - 0.852156i	0.989376
0.502688 + 0.852156i	0.989376
0.491220 - 0.857997i	0.988664
0.491220 + 0.857997i	0.988664
-0.987128	0.987128
0.841268 - 0.493908i	0.975539
0.841268 + 0.493908i	0.975539
0.012608 + 0.971084i	0.971166
0.012608 - 0.971084i	0.971166
0.962902 + 0.044752i	0.963941
0.962902 - 0.044752i	0.963941
-0.792333 + 0.498272i	0.935985
-0.792333 - 0.498272i	0.935985
0.929885 - 0.095356i	0.934761
0.929885 + 0.095356i	0.934761
-0.801131 - 0.467851i	0.927737
-0.801131 + 0.467851i	0.927737
-0.879480 - 0.246285i	0.913314
-0.879480 + 0.246285i	0.913314
-0.911202	0.911202
-0.540786 + 0.731983i	0.910082
-0.540786 - 0.731983i	0.910082
-0.433042 + 0.795654i	0.905865
-0.433042 - 0.795654i	0.905865
0.261812 - 0.863156i	0.901989
0.261812 + 0.863156i	0.901989
0.686826 - 0.560051i	0.886221
0.686826 + 0.560051i	0.886221
0.458032 - 0.733058i	0.864389
0.458032 + 0.733058i	0.864389
-0.643388 - 0.575394i	0.863149
-0.643388 + 0.575394i	0.863149
0.739740 + 0.444666i	0.863101
0.739740 - 0.444666i	0.863101
-0.072325 + 0.857117i	0.860163
-0.072325 - 0.857117i	0.860163
0.036855 - 0.856626i	0.857419
0.036855 + 0.856626i	0.857419
0.818844 - 0.100887i	0.825036
0.818844 + 0.100887i	0.825036
-0.802434 - 0.118723i	0.811169
-0.802434 + 0.118723i	0.811169
-0.227200 - 0.749577i	0.783253
-0.227200 + 0.749577i	0.783253
0.672394 + 0.321919i	0.745483
0.672394 - 0.321919i	0.745483
-0.043637	0.043637

VEC specification imposes 3 unit root(s).

Table 55: Bound tests for in-sample year-on-year unrestricted error correction forecasting model.

Wald Test:

Test Statistic	Value	df	Probability
F-statistic	9.734603	(4, 216)	0.0000
Chi-square	38.93841	4	0.0000

Null Hypothesis: $\beta_{L^1 p_{NL}} = \beta_{L^1 p_{EA}} = \beta_{L^1 u_{NL}} = \beta_{L^1 u_{EA}} = 0$

Table 56: Bound tests for in-sample annualized unrestricted error correction forecasting model.

Wald Test:

Test Statistic	Value	df	Probability
F-statistic	3.723150	(2, 206)	0.0258
Chi-square	7.446301	2	0.0242

Null Hypothesis: $\beta_{L^1 p_{NL}} = \beta_{L^1 p_{EA}} = 0$

M Appendix: Auxiliary Forecasts

Table 57: Summary of the ARIMA automatic model selection for the forecast of p_{EA}. Forecast sample: 01/1996-06/2011

Automatic ARIMA Forecasting
Selected dependent variable: Δp_{EA} (Out of p_{EA} and Δp_{EA}), Periodicity:12
Included observations: 185
Forecast length: 60

Number of estimated ARMA models: 169
Number of non-converged estimations: 0
Selected ARMA model: (10,9)
SIC value:-8.7708681117

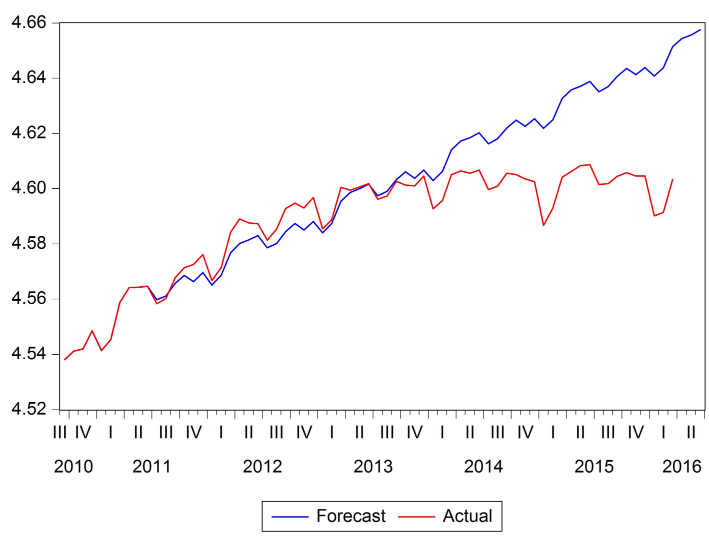


Figure 33: Actual versus the forecast of p_{EA}, 4 3/4 years.

Table 58: Summary of the ARIMA automatic model selection for the forecast of p_{EA}. Forecast sample: 01/1996-01/2015.

Automatic ARIMA Forecasting Selected dependent variable: Δp_{EA} (Out of p_{EA} and Δp_{EA}), Periodicity:12 Included observations: 228 Forecast length: 16
Number of estimated ARMA models: 169 Number of non-converged estimations: 1 Selected ARMA model: (12,11) SIC value: -11.8215036391

Figure 34: Actual versus the forecast of p_{EA}, 1 year.

Table 59: Summary of the ARIMA automatic model selection for the forecast of u_{NL}. Forecast sample: 01/1996-06/2011.

Automatic ARIMA Forecasting
Selected dependent variable: u_{NL} (Out of u_{NL} and Δu_{NL}), Periodicity:24
Note that the model includes additional non-linear terms.
Included observations: 183
Forecast length: 60

Number of estimated ARMA models: 225
Number of non-converged estimations: 0
Selected ARMA model: (3,2)
SIC value: -5.74213744122

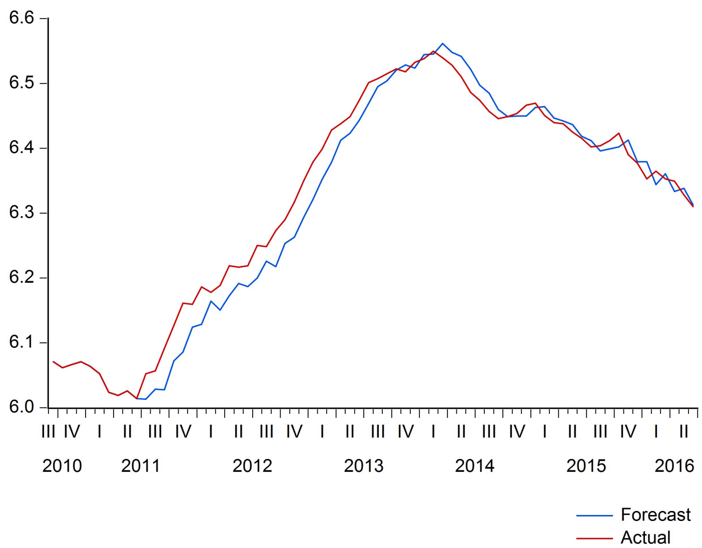


Figure 35: Actual versus the forecast of u_{NL}, 4 3/4 year.

Table 60: Summary of the ARIMA automatic model selection for the forecast of u_{NL}. Forecast sample: 01/1996-01/2015.

Automatic ARIMA Forecasting Selected dependent variable: Δu_{NL} (Out of u_{NL} and Δu_{NL}), Periodicity:24 Note that the model includes additional non-linear terms. Included observations: 226 Forecast length: 16
Number of estimated ARMA models: 169 Number of non-converged estimations: 0 Selected ARMA model: (4,3) SIC value: -5.78134995997

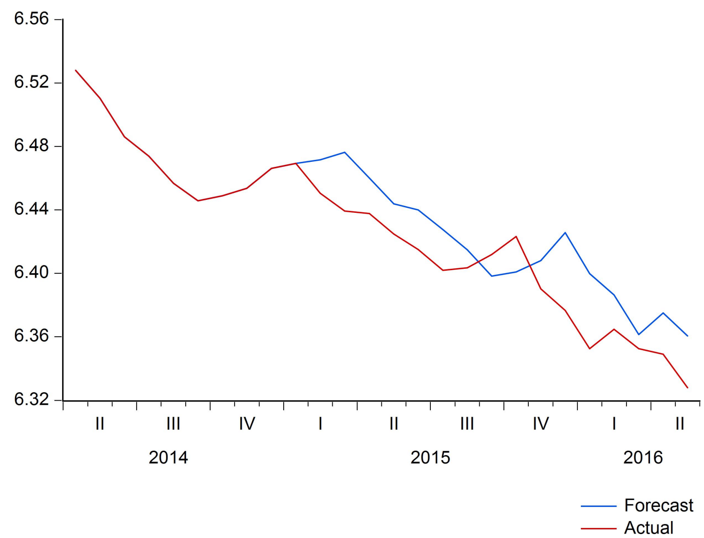


Figure 36: Actual versus the forecast of u_{NL}, 1 year.

Table 61: Summary of the ARIMA automatic model selection for the forecast of u_{EA}. Forecast sample: 01/1996-06/2011.

Automatic ARIMA Forecasting Selected dependent variable: u_{EA} (Out of u_{EA} and Δu_{EA}), Periodicity:24 Note that the model includes additional non-linear terms. Included observations: 184 Forecast length: 60
Number of estimated ARMA models: 25 Number of non-converged estimations: 0 Selected ARMA model: (3,0) SIC value: -6.8961809949

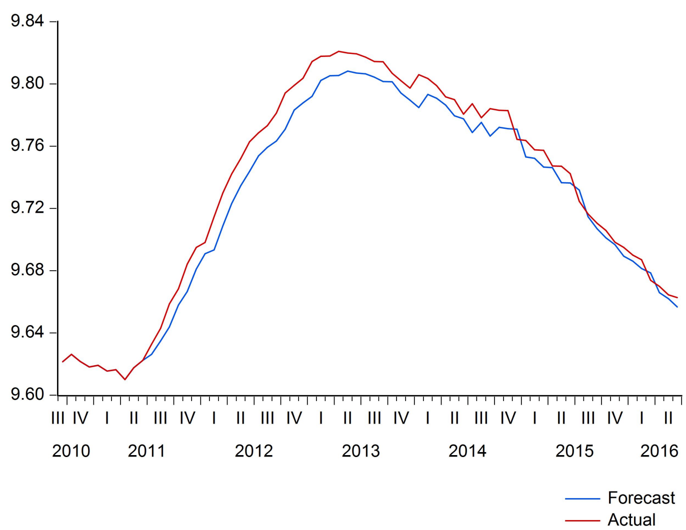


Figure 37: Actual versus the forecast of u_{EA}, 4 3/4 year.

Table 62: Summary of the ARIMA automatic model selection for the forecast of u_{EA}. Forecast sample: 01/1996-01/2015.

Automatic ARIMA Forecasting Selected dependent variable: u_{EA} (Out of Δu_{EA} and Δu_{EA}), Periodicity:24 Note that the model includes additional non-linear terms. Forecast length: 12
Number of estimated ARMA models: 1521 Number of non-converged estimations: 3 Selected ARMA model: (1,0) SIC value: -7.00839547411

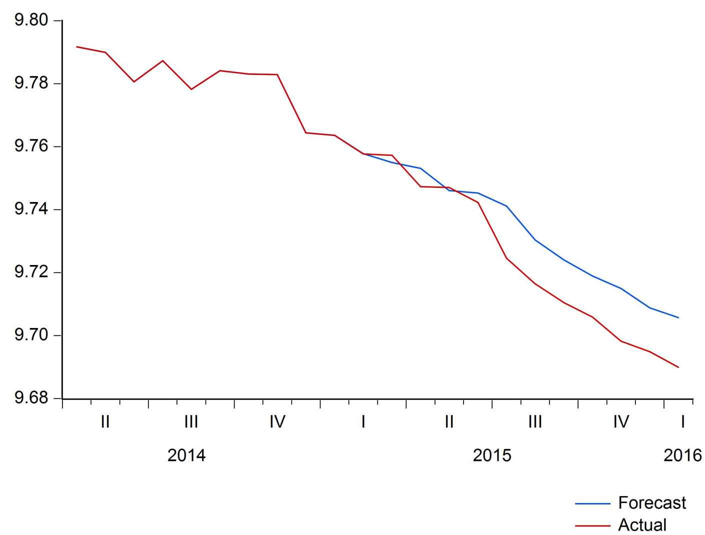


Figure 38: Actual versus the forecast of u_{EA}, 1 year.