

Norkute, Milda

Working Paper

Assessing the New Keynesian Phillips Curve in the Euro Area Using Disaggregate Data

Working Paper, No. 2013:31

Provided in Cooperation with:

Department of Economics, School of Economics and Management, Lund University

Suggested Citation: Norkute, Milda (2013) : Assessing the New Keynesian Phillips Curve in the Euro Area Using Disaggregate Data, Working Paper, No. 2013:31, Lund University, School of Economics and Management, Department of Economics, Lund

This Version is available at:

<https://hdl.handle.net/10419/260087>

Standard-Nutzungsbedingungen:

Die Dokumente auf EconStor dürfen zu eigenen wissenschaftlichen Zwecken und zum Privatgebrauch gespeichert und kopiert werden.

Sie dürfen die Dokumente nicht für öffentliche oder kommerzielle Zwecke vervielfältigen, öffentlich ausstellen, öffentlich zugänglich machen, vertreiben oder anderweitig nutzen.

Sofern die Verfasser die Dokumente unter Open-Content-Lizenzen (insbesondere CC-Lizenzen) zur Verfügung gestellt haben sollten, gelten abweichend von diesen Nutzungsbedingungen die in der dort genannten Lizenz gewährten Nutzungsrechte.

Terms of use:

Documents in EconStor may be saved and copied for your personal and scholarly purposes.

You are not to copy documents for public or commercial purposes, to exhibit the documents publicly, to make them publicly available on the internet, or to distribute or otherwise use the documents in public.

If the documents have been made available under an Open Content Licence (especially Creative Commons Licences), you may exercise further usage rights as specified in the indicated licence.

Working Paper 2013:31

Department of Economics
School of Economics and Management

Assessing the New Keynesian Phillips Curve in the Euro Area Using Disaggregate Data

Milda Norkute

September 2013



LUND
UNIVERSITY

Assessing the New Keynesian Phillips Curve in the Euro Area Using Disaggregate Data^{*}

Milda Norkute[†]

September 13, 2013

Abstract

There is no a priori reason to suppose that price-setting behaviour is homogeneous across sectors and countries. Aggregate data is, however, commonly used to estimate the New Keynesian Phillips curve (NKPC), which may very well yield erroneous results if price-setting behaviour is heterogeneous. In this paper we therefore estimate the hybrid NKPC for the Euro Area using a novel sectoral data set containing quarterly observations from 1999Q1 to 2012Q1. We show that a positive relationship between inflation and real marginal cost cannot be established empirically for a majority of countries and sectors. We also perform a meta-analysis by combining the results of individual significance tests in order to assess the validity of the NKPC in each country across all sectors and in each sector across all countries. We find no empirical evidence for the NKPC in the Euro Area when this meta-analysis is used. Our results therefore raise doubts about the appropriateness of the NKPC for the analysis of inflation dynamics and monetary policy in the Euro Area.

JEL classification: E31; C12

Keywords: New Keynesian Phillips curve; Inflation; Meta-analysis

^{*} I would like to thank Fredrik N. G. Andersson, David Edgerton and seminar participants at Lund University for constructive comments. Financial support from the Jan Wallander and Tom Hedelius Foundation is gratefully acknowledged.

[†] Corresponding address: Department of Economics, Lund University, P.O. Box 7082, SE-220 07 Lund, Sweden. Tel.: + 46 46 222 7919. E-mail address: Milda.Norkute@nek.lu.se.

1. Introduction

Empirical support for the New Keynesian Phillips curve (NKPC) is provided by, among others, Galí and Gertler (1999) and Sbordone (2002) for the United States and Galí *et al.* (2001) for the Euro Area. These studies use aggregate data, however, which has become the subject of criticism. Using aggregate data to estimate the NKPC will lead to biased parameter estimates if price-setting behaviour differs across sectors – the importance of real marginal cost will be underestimated and inflation persistence will be overestimated (see Imbs *et al.*, 2007). Along the same line of reasoning, it is plausible to assume that price-setting behaviour is not only heterogeneous across sectors but also across countries. This implies that estimating the NKPC for the Euro Area as a whole will yield biased parameter estimates.

Studies of sectoral NKPC for European countries have yielded inconclusive results. Substantial diversity in price-setting behaviour is reported by Imbs *et al.* (2007, 2011), who investigate disaggregate French data set and come to the conclusion that sectoral NKPC is supported on empirical grounds. On the other hand, Lawless and Whelan (2011) show that the NKPC does not provide a good approximation of sectoral inflation dynamics in the Euro Area. Mixed evidence is also provided in an extensive analysis by Byrne *et al.* (2013) using a disaggregate data set comprising several European countries and the United States –strong support is found for large economies (the U.S., the United Kingdom, Germany and France), but with contradicting results for smaller countries.

In this paper we assess the empirical validity of the New Keynesian Phillips curve for the Euro Area countries using a novel data set disaggregated by both country and sector. Previous research on the NKPC for European countries has relied on the EU KLEMS database. Only annual data is available in this database, however, and it has to be converted into quarterly data in order to estimate the structural parameters of the NKPC (see Leith and Malley, 2007). Our study uses a harmonised quarterly data set provided by Eurostat, however, and unlike previous studies our results are therefore not reliant on any statistical extrapolation method.

The NKPC is usually estimated by assuming that the rational expectations hypothesis holds and that inflation expectations can be replaced by actual inflation. In this case, however, the forward-looking component in the NKPC becomes correlated with the error term. In order to handle this endogeneity problem, the Generalized Method of Moments (GMM) estimator is typically used in the literature to estimate the NKPC (see among others Galí and Gertler, 1999; Galí *et al.*, 2001; Leith and Malley, 2007). In this study, we estimate the sectoral hybrid NKPC using Factor-GMM (see Bai and Ng, 2010; Kapetanios and Marcellino, 2010). Factor-

GMM is generally considered to be superior to GMM in terms of efficiency, since principal components that are estimated from a large data set of macroeconomic variables are used as additional instruments for expected inflation.

Finally, we carry out meta-analysis based on the combination of our results from the individual significance tests. In this way we can test the global null hypothesis that states that there is no relationship between real marginal cost and inflation in each country across all sectors and each sector across all countries. This meta-analysis is carried out using a Simes-type test (see Simes, 1986). This test is relatively new in the economic literature (see Hanck 2011, 2013), although it is widely used in genetics, pharmacology and other applied sciences (see Dmitrienko and Tamhane, 2007; Goeman and Solari, 2011). The key advantage of the Simes-type test is that it is valid even if the assumption of cross-sectional independence does not hold. This is important as the assumption of cross-sectional independence is unlikely to hold in our application, since idiosyncratic shocks are expected to be correlated both across sectors and countries due to prevalence of common factors and spatial dependence. We also implement Hommel's procedure (see Hommel, 1988) to identify the sectors and countries for which the individual hypotheses can be rejected.

Our analysis shows that the empirical support for the micro-founded NKPC is weak. In only two cases are the parameter estimates in line with economic theory, whereas the results are in direct contradiction to theory in most of the other cases. Furthermore, Hommel's procedure reveals that the parameter governing the link between inflation and the real marginal cost is statistically insignificant for the two cases which yield results in agreement with economic theory. This implies that the meta-analysis gives no empirical support for the NKPC in the Euro Area.

Changes in underlying market structure and technology within sectors of the Euro Area could account for the poor empirical evidence found in support of the micro-founded New Keynesian Phillips curve. A stable positive relationship between inflation and real marginal cost cannot therefore be expected if the structural parameters of the New Keynesian Phillips curve vary over time (see Lawless and Whelan, 2011). This raises the question of whether the theoretical framework used to conduct the monetary policy in the Euro Area should be revised, since the NKPC based on a Calvo-style sticky-price model is unable to explain both technological changes and shifts in economic conditions and market structures. The evaluation of alternative modelling frameworks for inflation dynamics in the Euro Area is beyond the scope of this paper, however.

The rest of paper is structured as follows. In section 2 the theoretical underpinnings are briefly outlined. The estimation method is discussed in section 3. The data set is described in section 4 and the empirical results are presented in section 5. Concluding remarks are found in section 6.

2. Theoretical Framework

Our theoretical model is based on Galí and Gertler (1999), Sbordone (2002) and Imbs *et al.* (2011). In this model each country $i = 1, \dots, I$ consists of sectors $j = 1, \dots, J$. Firms are assumed to be monopolistically competitive in the goods market and price rigidity is generated by introducing staggered price contracts (see Calvo, 1983). Furthermore, it is assumed that capital input is perfectly mobile across firms within each sector and country, implying that firms face the same marginal costs.

The sectoral version of the New Keynesian Phillips curve (NKPC) which links inflation with its expectations and real marginal cost is derived for country i and sector j as follows

$$\pi_{ij,t} = \beta_{ij} E_t \pi_{ij,t+1} + \frac{(1-\delta_{ij}\beta_{ij})(1-\delta_{ij})}{\delta_{ij}} s_{ij,t} + e_{ij,t}, \quad (1)$$

where inflation is denoted $\pi_{ij,t}$ and real marginal cost $s_{ij,t}$ is expressed as log-deviations from its respective long-run steady state. The term $e_{ij,t}$ represents a cost-push shock that accounts for other variation in inflation in addition to changes in excess demand (see Clarida, 1999; Woodford, 2003). The parameter β_{ij} is a subjective discount factor while δ_{ij} denotes the probability of firms not being able to set prices optimally. If prices are fully flexible ($\delta_{ij} \rightarrow 0$) firms are able to adjust prices and the slope coefficient associated with real marginal cost will be infinitely large. On the other hand, if there is complete price rigidity ($\delta_{ij} \rightarrow 1$) firms will set the same price as in the previous period and current inflation will not respond to the changes in real marginal cost. Different degrees of price rigidity can thus be accounted for as the probability δ_{ij} varies between zero and one. The average time over which the prices are constant will be given by $\tau_{ij} = 1/(1 - \delta_{ij})$ if the price changes are independent (see Galí and Gertler, 1999; Sbordone, 2002).

Following Galí and Gertler (1999), it is also assumed that a fraction of firms, denoted ω_{ij} , behave in a backward-looking fashion and set prices using previous changes in the price

index. The other firms make their pricing decisions in a forward-looking manner. The introduction of backward-looking firms leads to the following hybrid New Keynesian Phillips curve for country i and sector j ,

$$\pi_{ij,t} = \lambda_{ij}^b \pi_{ij,t-1} + \lambda_{ij}^f E_t \pi_{ij,t+1} + \varphi_{ij} s_{ij,t} + e_{ij,t}, \quad (2)$$

where the coefficients λ_{ij}^b , λ_{ij}^f and φ_{ij} are non-linear functions of the underlying structural parameters given by $\lambda_{ij}^b = \omega_{ij} \phi_{ij}^{-1}$, $\lambda_{ij}^f = \delta_{ij} \beta_{ij} \phi_{ij}^{-1}$ and $\varphi_{ij} = (1 - \omega_{ij})(1 - \delta_{ij})(1 - \delta_{ij} \beta_{ij}) \phi_{ij}^{-1}$ with $\phi_{ij} = \delta_{ij} + \omega_{ij}[1 - \delta_{ij}(1 - \beta_{ij})]$. Here, β_{ij} is a subjective discount factor, ω_{ij} is a fraction of backward-looking firms whereas δ_{ij} is a rigidity parameter or Calvo probability (see Calvo, 1983).

To be in line with economic theory the inequalities $\lambda_{ij}^b \geq 0$, $\lambda_{ij}^f > 0$ and $\varphi_{ij} > 0$ should hold (see Nymoen *et al.*, 2012). The hybrid NKPC reduces to the pure forward-looking New Keynesian Phillips curve given in Eq. (1) if the backwardness parameter λ_{ij}^b is equal to zero. We prefer the hybrid specification of NKPC since it is better capable of explaining the high degree of inflation persistence usually observed in actual data (see Fuhrer and Moore, 1995; Galí and Gertler, 1999).

Expected inflation can be replaced by actual inflation if we make the assumption of rational expectations, where economic agents do not make systematic mistakes when predicting future inflation. The amended version of the hybrid New Keynesian Phillips curve now takes the form

$$\pi_{ij,t} = \lambda_{ij}^b \pi_{ij,t-1} + \lambda_{ij}^f \pi_{ij,t+1} + \varphi_{ij} s_{ij,t} + \varepsilon_{ij,t}, \quad (3)$$

where the error term is defined as $\varepsilon_{ij,t} = e_{ij,t} - \lambda^f (\pi_{ij,t+1} - E_t \pi_{ij,t+1})$. The empirical version of the hybrid New Keynesian Phillips curve in Eq. (3) can thus be seen as a difference equation for inflation, where the rational expectations solution of Eq. (3) is given by

$$\pi_{ij,t} = \rho_{ij,1} \pi_{ij,t-1} + \varphi_{ij} (1 - \rho_{ij,1} \lambda_{ij}^f)^{-1} \sum_{s=0}^{\infty} \rho_{ij,2}^{-s} E_t s_{ij,t+s} + (1 - \rho_{ij,1} \lambda_{ij}^f)^{-1} \varepsilon_{ij,t}, \quad (4)$$

where $\rho_{ij,1}$ and $\rho_{ij,2}$ denote the stable and unstable roots of the characteristic equation¹ $P(x) = \lambda_{ij}^f x^2 - x + \lambda_{ij}^b = 0$. The rational expectation solution of the hybrid New Keynesian Phillips curve thus shows that present inflation is determined by lagged inflation, a discounted infinite sum of expected real marginal costs and a contemporaneous error term. It also shows that future real marginal costs are important in determining current inflation if prices are flexible (low δ_{ij} and high φ_{ij}). On the other hand expected future real marginal costs do not play an important role in inflation dynamics if the degree of price rigidity is high (high δ_{ij} and low φ_{ij}). In other words, a larger weight is assigned to future real marginal costs if a lower degree of price rigidity is present and vice versa.

3. Estimation method

The NKPC is commonly estimated using GMM (see among others Galí and Gertler, 1999; Galí *et al.*, 2001 and Leith and Malley, 2007) since the error term in Eq. (3) is correlated with inflation, which otherwise leads to an endogeneity problem. The appropriateness of using GMM estimation methods has, however, come under close scrutiny (see Lindè, 2005). Some disadvantages of the GMM estimator are poor small sample properties and problems posed by the use of weak instruments. These raise the need for an alternative estimation method to estimate the NKPC.

We estimate the sectoral hybrid NKPC in Eq. (3) using the Factor-GMM method, which has recently been proposed by Kapetanios and Marcellino (2010) and Bai and Ng (2010). The main advantage of the Factor-GMM estimation method is that it is more efficient than standard GMM. In particular, considerable gains in efficiency are obtained using Factor-GMM when the sample size is small and the instruments are weak (see Kapetanios and Marcellino, 2010). Beyer *et al.* (2008) has also shown that the standard errors for the coefficients associated with forward-looking variables are substantially smaller when Factor-GMM is employed instead of GMM.

The basic idea behind Factor-GMM is that a large number of potentially relevant instruments can be explained by a smaller number of factors that drive the data. If the endogenous variable is driven by the same exogenous factors as the observable instruments,

¹ The condition $\lambda_{ij}^f + \lambda_{ij}^b < 1$ is required to ensure that $|\rho_{ij,1}| < 1$ and $|\rho_{ij,2}| > 1$. In this case, a unique rational expectation solution exists and inflation is stationary. If $\lambda_{ij}^f + \lambda_{ij}^b = 1$ and $\lambda_{ij}^f < 0.5$, then $\rho_{ij,1} = 1$ implies that inflation is $I(1)$ since $s_{ij,t} \sim I(0)$ by construction. However, if $\lambda_{ij}^f + \lambda_{ij}^b > 1$, then a unique rational expectations solution does not exist (see Nymoen *et al.*, 2012 and Dees *et al.*, 2009).

then these factors will be valid and efficient instruments (see Bai and Ng, 2010). In the NKPC the endogenous variable is future inflation, which we assume depends on a few fundamental macroeconomic variables, inflation lag (due to persistence) and a large set of other economic variables that can be explained by few factors (see Kapetanios and Marcellino, 2010). The estimated common factors can thus be used as instruments for the forward-looking component in the New Keynesian Phillips curve. The unobserved factors can be consistently estimated by using principal components, which will allow for weak cross-sectional and serial correlation (see Stock and Watson, 2002). The number of factors is usually determined by minimising an information criterion (see Bai and Ng, 2002).

4. Data

Our data comes from the Eurostat and European Central Bank (ECB) databases. The data set includes all Euro Area countries except Portugal, Malta, Cyprus and Luxembourg, which are excluded due to missing observations. We use quarterly data where the sample starts in 1999Q1 and ends in 2012Q1. The starting point is motivated by the launch of the European Monetary Union (EMU) on the 1st of January in 1999.

We consider the sectors *Agriculture*, *Manufacturing*, *Construction*, *Services* and *Other Business Services* as given in NACE Rev. 2. The remaining sectors are not included in the analysis since they cover the public sector and a few very small branches². The description of sectors used in our analysis is provided in Appendix A, Table A.1.

The sector-specific variables are inflation and real marginal cost. Sector-level inflation is measured by the quarter-on-quarter growth rate of the value added deflator for the corresponding sector. We use a proxy for the sector-specific real marginal cost, namely the logarithm of the labour income share, which is computed as the ratio of compensation of employees to nominal value added in the respective sector. All the data is demeaned.

Some descriptive statistics for sector-specific inflation and labour shares for each country and sector are provided in Appendix A, Table A.2. and Table A.3. respectively. These show that the lowest average sector-specific inflation rate is in Germany, whereas the largest quarterly price changes are observed in the new Euro Area members i.e. Estonia, Slovakia and Slovenia. The largest average inflation rate is in *Construction*, but it is relatively modest in *Manufacturing* and *Agriculture*. The latter sector is also characterised by the largest standard

² Such as arts, entertainment, recreation, activities of households and extra-territorial organisations, etc.

deviation in both variables. The service sectors display low volatility compared to other sectors, however. In general the descriptive statistics indicate noticeable dissimilarities in inflation and labour share dynamics across countries and sectors, suggesting that a pronounced heterogeneity may be present at both sector and country level.

Following Galí and Gertler (1999), we use four external instruments – output gap, wage inflation, long-short interest rate spread and commodity price changes. The output gap is measured as a cyclical component obtained by filtering the real GDP with a Hodrick-Prescott filter³. Wage inflation is calculated as the quarterly growth rate of the labour cost index. The time series of 3-month Euribor and Euro Area 10-year government benchmark bond yield are used to construct the long-short interest rate spread. Finally, the non-energy commodity price index is used to compute commodity price inflation. A large set of macroeconomic variables is used to extract the country-specific factors that are employed as additional instruments. In total we use 27 variables. A description of the data that is used to estimate the factors is provided in Appendix A, Table A.4. The number of factors is based on the IC_2 information criterion⁴ (see Bai and Ng, 2002), which is allowed to be different for each country. The estimated number of factors is given in Appendix A, Table A.5.

Prior to estimation, all the data was corrected for outliers by applying the wavelet-based method described in Appendix B. Removing outliers is necessary since the standard principal components algorithm works poorly when any type of irregularities are present in the data. Our data set has to be denoised in this way since it is likely to contain extreme values caused by abrupt changes in macroeconomic conditions such as the recent financial crisis⁵. The instruments have also been demeaned and the data used to estimate factors has been standardised. All the series, except the unemployment rate, have been seasonally adjusted⁶.

The Bootstrap Sequential Quantile Test (BSQT) (see Smeekes, 2011) is employed to test whether or not the data is stationary. This test provides an estimate of the fraction of stationary units thus making possible to determine the stationarity properties of individual series in a panel. The results for the BSQT for sector-level inflation and labour shares are presented in Appendix A, Table A.6. The results show that all the series for sector-level inflation are stationary since the null hypothesis is rejected at the 5% level of significance for

³ The smoothing parameter for the Hodrick-Prescott filter is set to 1600, which is typically used for quarterly data.

⁴ The maximum number of factors is set to 8 times the smallest integer that is larger than $(T/100)^{0.25}$.

⁵ We have also estimated the sectoral hybrid NKPC using data that has not been corrected for outliers. We found that the results are rather sensitive to the presence of irregularities in the data. However, we found that even fewer cases yield statistically significant parameter estimates when outliers are not removed from the data, which is a conservative result when related to our findings.

⁶ The data is seasonally adjusted using a seasonal dummy variables regression.

all the countries. The BSQT test results indicate that labour shares are non-stationary for some countries, however. We disregard these unit root test results since following Dees *et al.* (2009) the marginal cost variable is stationary by construction. The BSQT test performed on the remaining variables show that most of the series are stationary with few exceptions⁷. The test results for the other variables are not reported, but they are available upon request.

Finally, the correlations between sector-specific inflation and the labour shares are reported in Appendix A, Table A.7. A positive dependence between the variables is only observed in a few cases, and this dependence seems to be rather weak since the correlations are small in magnitude. In most of countries and sectors the empirical correlations are in fact negative and relatively large in absolute value, which is the opposite of what is expected in economic theory. This already gives us some preliminary evidence to suspect that a positive relationship between inflation and real marginal cost is not supported by the data for a majority of sectors and countries.

5. Results

5.1. Estimation results

The estimation results for the sectoral hybrid New Keynesian Phillips curves for each sector and country are provided in Tables 1, 2 and 3. These contain the estimated forwardness and backwardness parameters λ_{ij}^f and λ_{ij}^b , for inflation lead and lag respectively, and the slope coefficient φ_{ij} associated with real marginal cost. The intercept is also included in the regressions, although estimation results of the constant term are not reported. The validity of model is tested by running the *J*-test for over-identifying restrictions, which shows that the null hypothesis of model validity is only rejected for *Services* in Estonia and *Other Business Services* in Greece. These results are available upon request.

The estimation results in Table 1 and 2 show that the forwardness and backwardness parameters are statistically significant in most of the cases. Some of them are significantly negative, however, which is in direct contradiction to economic theory. Moreover, no conclusion about dominant forward-looking behaviour could be drawn when comparing the

⁷ The BSQT method does not lead to a single rejection for unemployment and government expenditure as a ratio to GDP. These variables, however, are not excluded from the data set since according to economic theory they are considered to be mean reverting processes, therefore the failure to reject the null hypothesis of a unit root can be attributed to inefficiency due to small sample size.

magnitude of the forwardness and backwardness parameters. There is no clearly defined pattern showing that the coefficient on inflation lead tends to be higher than the parameter associated with inflation lag.

[Insert Table 1]

[Insert Table 2]

The estimated slope coefficients for real marginal cost in Table 3 show that only in a few cases are the results positive and statistically significant, namely for *Agriculture* in Belgium, Ireland and Italy, *Manufacturing* in Estonia, Germany and Ireland, and *Construction* in the Netherlands. On the other hand, there are several cases for which the estimated parameter for real marginal cost is statistically significant but negative, which does not accord with economic theory. There is not a single case for which the parameter for real marginal cost is statistically significant and has the correct sign for *Services* and *Other Business Services* in all the countries. In the vast majority of cases the parameter governing the link between inflation and the real marginal cost is statistically insignificant, indicating that there is no statistical evidence for any relationship between real marginal cost and inflation. It implies that in most of the sectors and member states of the Euro Area prices are not driven by real marginal cost and inflation does not react to changes in real economic activity.

[Insert Table 3]

Taking into account the estimation results of all parameters of the hybrid NKPC, the empirical support for sectoral hybrid New Keynesian Phillips curve is only provided for two cases – *Agriculture* in Belgium and *Manufacturing* in Germany. In these cases both the forwardness and backwardness parameters and the slope coefficient for real marginal cost are positive and statistically significant, as suggested by economic theory.

The estimates of structural parameters for *Agriculture* in Belgium and *Manufacturing* in Germany, as retrieved from the reduced form parameter estimates, are provided in Table 4 below.

[Insert Table 4]

Since the subjective discount factor is usually calibrated to slightly less than unity in the literature (see Galí and Gertler, 1999), the estimate for β_{ij} is considerably too small in both cases, corresponding to 0.282 for *Agriculture* in Belgium and 0.312 for *Manufacturing* in Germany. Moreover, the estimated proportion of backward-looking price setters is smaller for *Agriculture* in Belgium, implying that forward-looking behaviour of firms is more pronounced for *Manufacturing* in Germany. Finally, the estimated Calvo probabilities are 0.922 a *Manufacturing* in Germany and 0.983 for *Agriculture* in Belgium, which correspond

to an average time duration between price changes of 12.9 and 58.8 quarters respectively. This indicates a relatively high degree of price rigidity as the estimated probabilities are rather close to one. The structural parameter estimates for the other countries and sectors are not discussed since no empirical support for the hybrid New Keynesian Phillips curve is to be found for the remaining cases.

5.2. Meta-analysis

In addition to the individual significance tests for each country and sector, we want to investigate if there is a positive relationship between inflation and the labour share in any country or sector in the Euro Area. In other words, we want to see if the hybrid New Keynesian Phillips curve can be reconciled with the data for some, but not all, of the Euro Area countries or sectors.

In order to identify those countries and sectors for which the NKPC could be judged valid on empirical grounds, we employ a meta-analytic approach which is based on combining the individual test results. We therefore test the global null hypothesis that the coefficient of real marginal cost is zero for a given country (or sector) across all sectors (or countries). To do this we use a Simes-type test, which is based on a modified Bonferroni multiple testing procedure (see Simes, 1986) usually referred in the literature as Simes' classical intersection test. An attractive feature of the Simes-type test is that it is valid even when individual tests are correlated, thus allowing for the presence of arbitrary cross-sectional dependence (see Hanck, 2013). It is also a relatively simple test based on a combination of the p -values for the individual tests.

The basic idea of the Simes-type test is that the p -values of the individual tests are ordered from smallest to largest $p_{(1)} \leq p_{(2)} \leq \dots \leq p_{(l)} \leq \dots \leq p_{(N)}$ and then compared with gradually increasing critical points (see Simes, 1986). The ordered p -values are denoted by $p_{(l)}$ where l is the position of p -value in the ordered series. The global null hypothesis is rejected if

$$p_{(l)} \leq \alpha l / N \quad \text{for some } l, \quad (5)$$

where α is the overall significance level. If the inequality in (5) holds, and the global null hypothesis is thus rejected, the question is raised as to which of individual hypotheses can be rejected. A multiple testing procedure is employed for this purpose, which controls the

Familywise Error Rate⁸ (FWER) so that this never exceeds a prespecified overall significance level α . Controlling FWER is necessary in order to avoid the multiplicity problem, which results in overstating the number of individual rejections. Hommel's procedure is defined as follows (see Hommel, 1988)

- i. compute $m = \max \{n \in \{1, \dots, N\}: p_{(N-n+k)} > k\alpha/n \text{ for } k \in \{1, \dots, n\}\}$;
- ii. if the maximum does not exist, reject all individual null hypotheses for $n = 1, \dots, N$;
- iii. otherwise, reject all individual null hypotheses with p -values that are smaller or equal to α/m .

In our analysis the p -values ordered across sectors and countries, with the corresponding critical points, are shown in Tables 5 and 6 respectively. The results in Table 5 show that the p -values which are smaller than the critical points corresponding to the 1% level of significance are for *Construction* in Estonia, *Services* in France, *Agriculture* and *Construction* in Germany, *Agriculture* in Italy and *Services* in Slovenia. Hence, at the 1% level of significance the global null hypothesis is rejected for Estonia, France, Germany, Italy and Slovenia as there is at least one sector for which the inequality defined in (5) holds. At the 5% level of significance the global null hypothesis can also be rejected for Belgium and Ireland, since the p -values for *Manufacturing* in Belgium and *Agriculture* and *Services* in Ireland are smaller than the respective critical points.

[Insert Table 5]

When p -values are combined across countries, the Simes-type test results in Table 6 show that the global null hypothesis stating that the slope coefficient for real marginal cost is jointly equal to zero across all countries can be rejected at the 1% level of significance for *Agriculture*, *Construction* and *Services*. At the 5% level of significance the global null hypothesis can also be rejected for *Manufacturing*. The global hypothesis cannot be rejected for *Other Business Services* at any conventional level of statistical significance⁹.

⁸ In each individual test, a probability to reject erroneously the null hypothesis corresponds to significance level α . As an example, a probability for at least one false rejection in n individual independent tests is $1 - (1 - \alpha)^n$. In this case, at the 5% significance level, the FWER are $1 - (1 - 0.05)^{13} = 0.487$ and $1 - (1 - 0.05)^5 = 0.226$ when joint significance is tested across all sectors and countries respectively, which are obviously larger than the desired overall significance level of 0.05.

⁹ Fisher-type tests (see Fisher, 1932) have also been performed as a robustness check. These yield similar results to the Simes-type test. The Fisher-type test results for each country or sector can be found in Appendix A, Tables A.8 and A.9 respectively. In addition, the algorithm used to obtain the critical values for the Fisher-type test is presented in Appendix C. Using Fisher-type tests, the global null hypothesis can be rejected for Estonia at the 1% level of statistical significance and for Belgium, France, Ireland, Slovenia and Spain at the 5% level of

[Insert Table 6]

The individual hypotheses rejected by Hommel's procedure are provided in Tables 7 and 8 for sectors and countries respectively. The results at the 5% level of significance are also presented graphically in Fig. 1 and Fig. 2, which show the ordered p -values (depicted as black dots) and the critical points (depicted as straight lines). The individual hypotheses have to be rejected when their p -values lie below the horizontal dashed line.

[Insert Table 7]

[Insert Table 8]

The results in Table 7 and Table 8 show that the parameter estimates are all negative for those cases where the individual null hypotheses are rejected using Hommel's procedure, with the exception of *Agriculture* in Ireland and Italy. In these latter two cases, however, the parameter associated with inflation lead is negative, which is also at odds with economic theory. Most importantly, for the two cases which yield the parameter estimates consistent with theory, namely, *Agriculture* in Belgium and *Manufacturing* in Germany, the individual null hypotheses are not rejected when the multiple testing procedure is performed.

All the individual hypotheses which are rejected using Hommel's procedure thus correspond to cases for which our other results do not support the sectoral hybrid NKPC. It implies that no empirical support for the NKPC can be found when a meta-analytic approach to testing is implemented.

6. Conclusions

In this paper we investigate whether or not the New Keynesian Phillips curve is supported empirically in the Euro Area. For this purpose, we estimate the sectoral hybrid NKPC using a novel disaggregate data set. We employ the Factor-GMM estimation method which is in general more efficient than usual GMM. The study also uses meta-analytic approach to combine the results of the individual significance tests. In this way, we can test the global null hypotheses that the parameter governing the relationship between inflation and real marginal cost is equal to zero for each country across all sectors and for each sector across all countries.

statistical significance when p -values are combined across sectors. When p -values are combined across countries the global null hypothesis cannot be rejected for *Agriculture* and *Other Business Services* at any conventional level of statistical significance. Hence, the Fisher-type test results are in agreement with the Simes-type test results. It should be noted, however, that the results from the Simes-type test are not completely comparable with the conclusions drawn using Fisher-type test, since in the latter case a different set of instruments has been used in the estimation. A few cases were also eliminated when performing the Fisher-type tests, whereas all countries and sectors are taken into account when implementing the Simes-type tests.

Based on the individual tests results we can conclude that the parameter associated with real marginal cost is statistically insignificant in most cases. For a few countries and sectors statistically significant negative parameter estimates are obtained, which also does not support the NKPC since economic theory implies a positive relationship between inflation and real marginal cost. The results are only in agreement with the NKPC in two cases, namely for *Agriculture* in Belgium and *Manufacturing* in Germany. The empirical support for NKPC is thus very weak, since the data is not reconciled with theory in the vast majority of cases.

The results of the meta-analysis show that the global null hypothesis is rejected at the 5% level of significance for Belgium, Estonia, France, Germany, Ireland, Italy and Slovenia when the p -values are combined across sectors using the Simes-type test. The global null hypothesis is also rejected for all sectors except *Other Business Services* when p -values are combined across countries. However, Hommel's procedure shows that these significant results are mainly due to the presence of *negative* parameter estimates, which is at odds with economic theory. In the only two cases which yield parameter estimates supporting the theory (*i.e.*, *Agriculture* in Belgium and *Manufacturing* in Germany), the individual hypotheses cannot be rejected by applying Hommel's procedure. The meta-analysis thus shows no evidence in favour of the NKPC. Our findings therefore cast doubt on the empirical validity of the New Keynesian Phillips curve in the Euro Area.

The failure to find statistical evidence for the NKPC can maybe be attributed to the relatively restrictive assumptions of the Calvo model of sticky prices, such as the absence of technological change or time-invariant market structures within sectors (see Lawless and Whelan, 2011). An assessment of alternative theoretical frameworks for inflation dynamics in the Euro Area is therefore an interesting topic for future research.

References

1. Bai, J., Ng, S., 2002. Determining the number of factors in approximate factor models. *Econometrica* 70 (1), 191-221.
2. Bai, J., Ng, S., 2010. Instrumental variable estimation in a data rich environment. *Econometric Theory* 26 (6), 1577-1606.
3. Bilen, C., Huzurbazar, S., 2002. Wavelet-Based Detection of Outliers in Time Series. *Journal of Computational and Graphical Statistics* 11 (2), 311-327.
4. Beyer, A., Farmer, R. E. A., Henry, J., Marcellino, M., 2008. Factor analysis in a model with rational expectations. *Econometrics Journal* 11 (2), 271-286.
5. Byrne, J. P., Kontonikas, A., Montagnoli A., 2013. International evidence on the New Keynesian Phillips curve using aggregate and disaggregate data. *Journal of Money, Credit and Banking* 45 (5), 913–932.
6. Calvo, G. A., 1983. Staggered prices in a utility-maximizing framework. *Journal of Monetary Economics*, 12 (3): 383–398.
7. Clarida, R., Galí, J., Gertler, M., 1999. The science of monetary policy: a new Keynesian perspective. *Journal of Economic Literature* 37 (4), 1661-1707.
8. Dees, S., Pesaran, M. H., Smith L. V., Smith R. P., 2009. Identification of New Keynesian Phillips curves from a global perspective. *Journal of Money, Credit and Banking* 41 (7), 1481-1502.
9. Dmitrienko, A., Tamhane, A. C., 2007. Gatekeeping procedures with clinical trial applications. *Pharmaceutical Statistics* 6 (3), 171-180.
10. Fisher, R. A., 1932. *Statistical methods for research workers*. Edinburgh, 4th Edition.
11. Fuhrer, J., Moore, G., 1995. Inflation persistence. *The Quarterly Journal of Economics* 110 (4), 127–159.
12. Galí, J., Gertler, M., 1999. Inflation dynamics: a structural econometric analysis. *Journal of Monetary Economics* 44 (2), 195-222.
13. Galí, J., Gertler, M., López-Salido, J. D., 2001. European inflation dynamics. *European Economic Review* 45 (7), 1237-1270.
14. Gençay, R., Selçuk, F., Whitcher, B., 2002. *An introduction to wavelets and other filtering methods in finance and economics*. Economic Press, San Diego.

15. Goeman, J. J., Solari, A., 2011. Multiple Testing for Exploratory Research. *Statistical Science* 26 (4), 584-597.
16. Granè, A., Veiga, H., 2010. Wavelet-based detection of outliers in financial time series. *Computational Statistics and Data Analysis* 54 (11), 2580-2593.
17. Hanck, C., 2011. Now, whose schools are really better (or weaker) than Germany's? A multiple testing approach. *Economic Modelling* 28 (4), 1739-1746.
18. Hanck, C., 2013. An intersection test for panel unit roots. *Econometric Reviews* 32 (2), 183-203.
19. Harvey, D. I., Leybourne, S. J., Taylor, R. A. M., 2012. Testing for unit roots in the presence of uncertainty over both the trend and initial condition. *Journal of Econometrics* 169 (2), 188–195.
20. Hommel, G., 1988. A stagewise rejective multiple tests procedure based on a modified bonferroni test. *Biometrika* 75 (2), 383–386.
21. Im, K. S., Pesaran, M. H., Shin, Y., 2003. Testing for unit roots in heterogeneous panels. *Journal of Econometrics* 115 (1), 53–74.
22. Imbs, J., Jondeau, E., Pelgrin, F., 2007. Aggregating Phillips curves. ECB Working Paper, No. 785.
23. Imbs, J., Jondeau, E., Pelgrin, F., 2011. Sectoral Phillips curves and the aggregate Phillips curve. *Journal of Monetary Economics* 58 (4), 328–344.
24. Inoue, A., Shintani, M., 2006. Bootstrapping GMM estimators for time series. *Journal of Econometrics* 133 (2), 531–555.
25. Kapetanios, G., Marcellino, M., 2010. Factor-GMM estimation with large sets of possibly weak instruments. *Computational Statistics and Data Analysis* 54 (11), 2655-2675.
26. Lawless, M., Whelan, K. T., 2011. Understanding the dynamics of labor shares and inflation. *Journal of Macroeconomics* 33 (2), 121–136.
27. Leith, C., Malley, J., 2007. A sectoral analysis of price-setting behavior in U.S. manufacturing industries. *Review of Economics and Statistics* 89 (2), 335–342.
28. Lindè, J., 2005. Estimating New-Keynesian Phillips curves: a full information maximum likelihood approach. *Journal of Monetary Economics* 52 (6), 1135–1149.

29. Nymoen, R., Swensen, A. R., Tveter, E., 2012. Interpreting the evidence for New Keynesian models of inflation dynamics. *Journal of Macroeconomics* 34 (2), 253–263.
30. Sbordone, A. M., 2002. Prices and unit labor costs: a new test of price stickiness. *Journal of Monetary Economics* 49 (2), 265–292.
31. Simes, R. J., 1986. An improved Bonferroni procedure for multiple tests of significance. *Biometrika* 73 (3), 751–754.
32. Smeekes, S., 2011. Bootstrap sequential tests to determine the stationary units in a panel. METEOR Working Paper RM/11/003.
33. Stock, J. H., Watson, M. W., 2002. Forecasting using principal components from a large number of predictors. *Journal of the American Statistical Association* 97 (460), 1167–1179.
34. Woodford, M., 2003. *Interest and Prices*. Princeton University Press.

Tables and figures

Table 1

Estimation results of the forwardness parameter λ_{ij}^f

	<i>Agriculture</i>	<i>Manufacturing</i>	<i>Construction</i>	<i>Services</i>	<i>Other Business Services</i>
Austria	0.466** (0.219)	0.297*** (0.077)	0.372*** (0.093)	0.378*** (0.095)	0.659*** (0.075)
Belgium	0.238*** (0.086)	0.323*** (0.071)	0.825*** (0.104)	0.475*** (0.05)	0.221 (0.166)
Estonia	0.367*** (0.129)	-0.044 (0.076)	-0.165*** (0.056)	-0.356 (1.084)	0.129 (0.14)
Finland	-0.087 (0.221)	0.208*** (0.063)	0.58*** (0.026)	-0.99* (0.578)	0.495 (0.298)
France	0.591*** (0.028)	0.408*** (0.12)	0.431** (0.2)	0.548*** (0.119)	0.257 (0.179)
Germany	0.109* (0.061)	0.28*** (0.065)	-0.145** (0.07)	-0.1 (0.1)	0.387*** (0.096)
Greece	-0.501 (0.323)	0.307 (0.221)	-0.758*** (0.123)	0.39*** (0.082)	2.128** (1.032)
Ireland	-0.365*** (0.101)	-0.453** (0.204)	0.232*** (0.056)	-0.151 (0.166)	0.413* (0.224)
Italy	-0.212 (0.153)	0.658*** (0.062)	-0.247* (0.144)	0.293*** (0.058)	0.079 (0.306)
Netherlands	0.101 (0.238)	0.067 (0.111)	0.104 (0.072)	-0.133 (0.081)	0.436 (0.374)
Slovakia	-0.09 (0.185)	0.311 (0.303)	0.631 (0.454)	-0.159 (0.153)	-0.473 (0.438)
Slovenia	0.116 (0.13)	0.436*** (0.034)	0.585*** (0.127)	-0.028 (0.059)	-0.286** (0.114)
Spain	-0.525*** (0.182)	-0.098 (0.085)	0.159 (0.107)	0.007 (0.141)	0.159 (0.134)

Notes: *,** and *** indicate the statistical significance at the 10%, 5% and 1% levels respectively. The numbers in the parentheses are standard errors corrected with HAC estimator (Bartlett spectral density kernel is used and the bandwidth is equal to 4). In line with Bayer *et al.* (2008), the instrument set contains the first lag of estimated factors. In addition, the set of instruments is augmented by four observable variables i.e. output gap, wage inflation, long-short interest rate spread, commodity price changes and their lags thereof (up to the 5th lag). Moreover, the lags of sectoral inflation and sector-specific labour shares are also used as instruments (up to the 5th lag). The set of instruments is reduced further by taking the most relevant ones using a hard thresholding method. In this way, the instruments are preselected by choosing only those for which *t* statistic exceeds unity in the first stage regression. In the estimation, we allow the set of instruments to be distinct in each case.

Table 2Estimation results of the backwardness parameter λ_{ij}^b

	<i>Agriculture</i>	<i>Manufacturing</i>	<i>Construction</i>	<i>Services</i>	<i>Other Business Services</i>
Austria	0.322 (0.556)	0.23 (0.155)	0.416*** (0.074)	0.081 (0.169)	0.42*** (0.083)
Belgium	0.463*** (0.087)	-0.022 (0.112)	0.29*** (0.077)	0.417*** (0.082)	0.383*** (0.124)
Estonia	0.37*** (0.079)	0.585*** (0.136)	0.042 (0.106)	-0.069 (1.154)	0.301** (0.117)
Finland	-0.138 (0.168)	0.463*** (0.083)	0.487*** (0.023)	1.989*** (0.697)	0.29* (0.163)
France	0.583*** (0.027)	0.529*** (0.135)	0.609*** (0.17)	-0.213** (0.098)	0.689*** (0.177)
Germany	0.087 (0.053)	0.318*** (0.041)	0.166 (0.148)	0.053 (0.092)	0.32** (0.128)
Greece	-0.444 (0.271)	0.461* (0.257)	0.157 (0.259)	0.403*** (0.071)	1.387 (1.006)
Ireland	-0.154* (0.085)	-0.302* (0.163)	0.389*** (0.103)	-0.485** (0.195)	0.261 (0.246)
Italy	0.318 (0.235)	-0.022 (0.095)	0.196 (0.242)	0.464*** (0.061)	0.199 (0.815)
Netherlands	-0.069 (0.172)	0.175* (0.088)	0.501*** (0.086)	-0.623*** (0.083)	0.761* (0.413)
Slovakia	0.021 (0.097)	0.143 (0.459)	1.18** (0.507)	-0.436** (0.175)	0.019 (1.084)
Slovenia	0.176 (0.137)	0.127 (0.109)	0.589 (0.395)	-0.596*** (0.059)	-0.205* (0.108)
Spain	-0.565** (0.251)	-0.084 (0.136)	0.172*** (0.055)	-0.055 (0.148)	0.195 (0.185)

Notes: *, ** and *** indicate the statistical significance at the 10%, 5% and 1% levels respectively. The numbers in parentheses are standard errors corrected with the HAC estimator (the Bartlett spectral density kernel is used and the bandwidth is equal to 4). In line with Bayer *et al.* (2008), the instrument set contains the first lag of the estimated factors. In addition, the set of instruments is augmented by four observable variables, *i.e.*, output gap, wage inflation, long-short interest rate spread, commodity price changes and their lags (up to the 5th lag). Moreover, the lags of sectoral inflation and sector-specific labour shares are also used as instruments (up to the 5th lag). The set of instruments is reduced by only using the most relevant ones using a hard thresholding method. In this way, the instruments are preselected by choosing only those for which *t* statistic exceeds unity in the first stage regression. We allow the set of instruments to be distinct in each case country and sector.

Table 3Estimation results of the coefficient associated with real marginal cost φ_{ij}

	<i>Agriculture</i>	<i>Manufacturing</i>	<i>Construction</i>	<i>Services</i>	<i>Other Business Services</i>
Austria	0.037 (0.065)	-0.006 (0.007)	-0.024** (0.012)	-0.063* (0.032)	-0.007 (0.011)
Belgium	0.026* (0.014)	-0.061*** (0.021)	0.003 (0.014)	-0.006 (0.005)	0.008 (0.021)
Estonia	-0.055 (0.05)	0.051* (0.03)	-0.169*** (0.026)	-0.059 (0.053)	0.00 (0.012)
Finland	0.144 (0.169)	0.002 (0.009)	0.01 (0.006)	0.009 (0.035)	-0.002 (0.016)
France	0.006 (0.008)	0.011 (0.017)	0.004 (0.024)	-0.158*** (0.028)	0.007 (0.035)
Germany	-0.154*** (0.029)	0.007* (0.004)	-0.046*** (0.014)	-0.007 (0.02)	0.019 (0.027)
Greece	-0.049 (0.041)	-0.013 (0.036)	0.006 (0.024)	0.012 (0.013)	-0.474* (0.27)
Ireland	0.074*** (0.025)	0.079* (0.04)	-0.054** (0.025)	-0.168** (0.064)	-0.12 (0.08)
Italy	0.109*** (0.032)	0.019 (0.017)	-0.033 (0.03)	0.003 (0.009)	-0.137 (0.091)
Netherlands	-0.149* (0.074)	-0.049** (0.02)	0.024* (0.014)	-0.002 (0.029)	-0.003 (0.065)
Slovakia	0.032 (0.033)	-0.005 (0.035)	0.022 (0.056)	-0.266** (0.112)	-0.272 (0.201)
Slovenia	-0.013 (0.031)	-0.038*** (0.013)	-0.002 (0.015)	-0.296*** (0.043)	-0.2** (0.078)
Spain	-0.08* (0.041)	-0.03** (0.013)	0.017 (0.013)	0.017 (0.034)	-0.201 (0.144)

Notes: *, ** and *** indicate the statistical significance at the 10%, 5% and 1% levels respectively. The numbers in parentheses are standard errors corrected with the HAC estimator (the Bartlett spectral density kernel is used and the bandwidth is equal to 4). In line with Bayer *et al.* (2008), the instrument set contains the first lag of the estimated factors. In addition, the set of instruments is augmented by four observable variables, *i.e.*, output gap, wage inflation, long-short interest rate spread, commodity price changes and their lags (up to the 5th lag). Moreover, the lags of sectoral inflation and sector-specific labour shares are also used as instruments (up to the 5th lag). The set of instruments is reduced by only using the most relevant ones using a hard thresholding method. In this way, the instruments are preselected by choosing only those for which *t* statistic exceeds unity in the first stage regression. We allow the set of instruments to be distinct in each case country and sector.

Table 4

Estimates of structural parameters

	Belgium	Germany
	<i>Agriculture</i>	<i>Manufacturing</i>
Subjective discount factor, β_{ij}	0.282	0.312
Fraction of backward-looking price setters, ω_{ij}	0.506	0.348
Rigidity parameter, δ_{ij}	0.922	0.983
Average time between price changes, τ_{ij}	12.851	58.866

Table 5The p -values of individual significance tests for φ_{ij} ordered across sectors

<i>Order of p-values</i>	<i>1</i>	<i>2</i>	<i>3</i>	<i>4</i>	<i>5</i>
Critical values ($\alpha = 0.01$)	0.002	0.004	0.006	0.008	0.010
Critical values ($\alpha = 0.05$)	0.01	0.02	0.03	0.04	0.05
Critical values ($\alpha = 0.1$)	0.02	0.04	0.06	0.08	0.10
Austria	(F) 0.047	(GI) 0.059	(BE) 0.409	(JN) 0.492	(A) 0.575
Belgium	(BE) 0.005	(A) 0.066	(GI) 0.269	(JN) 0.693	(F) 0.813
Estonia	(F) 0.000	(BE) 0.100	(A) 0.277	(GI) 0.279	(JN) 0.979
Finland	(F) 0.105	(A) 0.399	(GI) 0.799	(BE) 0.811	(JN) 0.910
France	(GI) 0.000	(A) 0.486	(BE) 0.524	(JN) 0.838	(F) 0.871
Germany	(A) 0.000	(F) 0.003	(BE) 0.075	(JN) 0.483	(GI) 0.723
Greece	(JN) 0.087	(A) 0.237	(GI) 0.380	(BE) 0.715	(F) 0.817
Ireland	(A) 0.006	(GI) 0.012	(F) 0.039	(BE) 0.055	(JN) 0.141
Italy	(A) 0.001	(JN) 0.140	(BE) 0.266	(F) 0.275	(GI) 0.720
Netherlands	(BE) 0.020	(A) 0.051	(F) 0.085	(GI) 0.957	(JN) 0.957
Slovakia	(GI) 0.022	(JN) 0.185	(A) 0.336	(F) 0.688	(BE) 0.890
Slovenia	(GI) 0.000	(BE) 0.007	(JN) 0.015	(A) 0.675	(F) 0.910
Spain	(BE) 0.028	(A) 0.058	(JN) 0.170	(F) 0.192	(GI) 0.616

Notes: order 1 indicates the smallest p -value and 5 denotes the largest one across all sectors in a corresponding country. Here, a shorthand notation for sectors is used: (A) - Agriculture, (BE) - Manufacturing, (F) - Construction, (GI) - Services, (JN) - Other Business Services.

Table 6

The p -values of individual significance tests for φ_{ij} ordered across sectors

<i>Order of p-values</i>	<i>1</i>	<i>2</i>	<i>3</i>	<i>4</i>	<i>5</i>	<i>6</i>	<i>7</i>	<i>8</i>	<i>9</i>	<i>10</i>	<i>11</i>	<i>12</i>	<i>13</i>
Critical values ($\alpha = 0.01$)	0.001	0.002	0.002	0.003	0.004	0.005	0.005	0.006	0.007	0.008	0.008	0.009	0.010
Critical values ($\alpha = 0.05$)	0.004	0.008	0.012	0.015	0.019	0.023	0.027	0.031	0.035	0.038	0.042	0.046	0.050
Critical values ($\alpha = 0.1$)	0.008	0.015	0.023	0.031	0.038	0.046	0.054	0.062	0.069	0.077	0.085	0.092	0.100
<i>Agriculture</i>	GE	IT	IR	NE	SP	BE	GR	ES	SK	FI	FR	AU	SL
	0.000	0.001	0.006	0.051	0.058	0.066	0.237	0.277	0.336	0.399	0.486	0.575	0.675
<i>Manufacturing</i>	BE	SL	NE	SP	IR	GE	ES	IT	AU	FR	GR	FI	SK
	0.005	0.007	0.020	0.028	0.055	0.075	0.100	0.266	0.409	0.524	0.715	0.811	0.890
<i>Construction</i>	ES	GE	IR	AU	NE	FI	SP	IT	SK	BE	GR	FR	SL
	0.000	0.003	0.039	0.047	0.085	0.105	0.192	0.275	0.688	0.813	0.817	0.871	0.910
<i>Services</i>	SL	FR	IR	SK	AU	BE	ES	GR	SP	IT	GE	FI	NE
	0.000	0.000	0.012	0.022	0.059	0.269	0.279	0.380	0.616	0.720	0.723	0.799	0.957
<i>Other Business Services</i>	SL	GR	IT	IR	SP	SK	GE	AU	BE	FR	FI	NE	ES
	0.015	0.087	0.140	0.141	0.170	0.185	0.483	0.492	0.693	0.838	0.910	0.957	0.979

Notes: order 1 indicates the smallest p -value and 13 denotes the largest one across all countries in a corresponding sector. Here, a shorthand notation for countries is used: BE – Belgium, ES – Estonia, FI – Finland, FR – France, GE – Germany, GR – Greece, IR – Ireland, IT – Italy, NE – Netherlands, SK – Slovakia, SN – Slovenia, SP – Spain.

Table 7Results from Hommel's procedure by combining p -values across sectors

	<i>Agriculture</i>	<i>Manufacturing</i>	<i>Construction</i>	<i>Services</i>	<i>Other Business Services</i>
Austria					
Belgium		(-)**			
Estonia			(-)**		
Finland					
France				(-)**	
Germany	(-)**		(-)**		
Greece					
Ireland	(+)**		(-)*	(-)**	
Italy	(+)**				
Netherlands					
Slovakia					
Slovenia		(-)**		(-)**	(-)**
Spain					

Notes: *,** and *** indicate statistical significance at the 10%, 5% and 1% levels respectively. - / + in the parentheses correspond to the negative/ positive sign of the parameter estimate for φ_{ij} .

Table 8Results from Hommel's procedure by combining p -values across countries

	<i>Agriculture</i>	<i>Manufacturing</i>	<i>Construction</i>	<i>Services</i>	<i>Other Business Services</i>
Austria					
Belgium		(-)*			
Estonia			(-)**		
Finland					
France				(-)**	
Germany	(-)**		(-)**		
Greece					
Ireland	(+)*				
Italy	(+)**				
Netherlands					
Slovakia					
Slovenia		(-)*		(-)**	
Spain					

Notes: *,** and *** indicate statistical significance at the 10%, 5% and 1% levels respectively. - / + in the parentheses correspond to the negative/ positive sign of the parameter estimate for φ_{ij} .

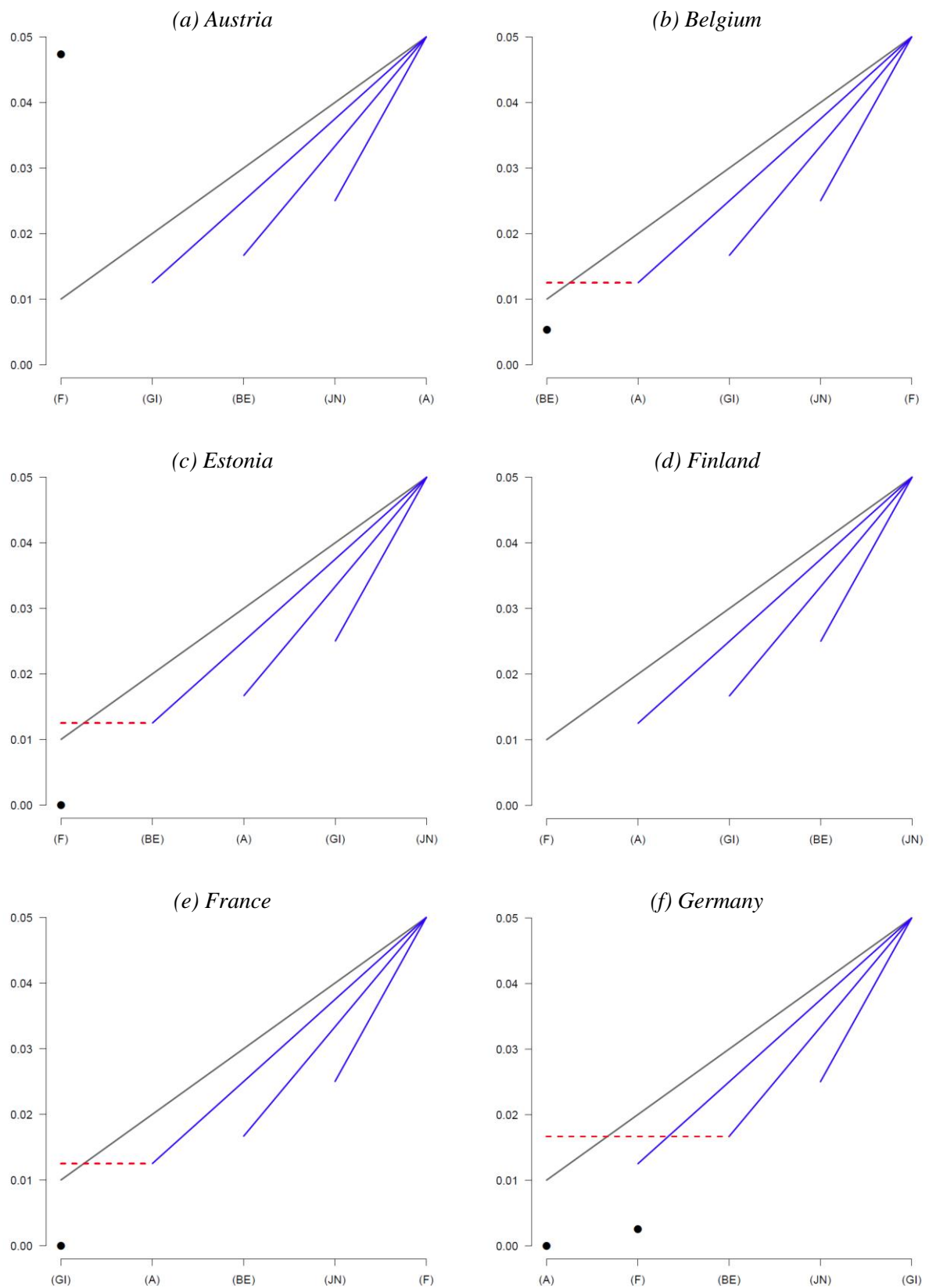


Fig. 1. Hommel's procedure by combining p -values across sectors (at the 5% significance level)

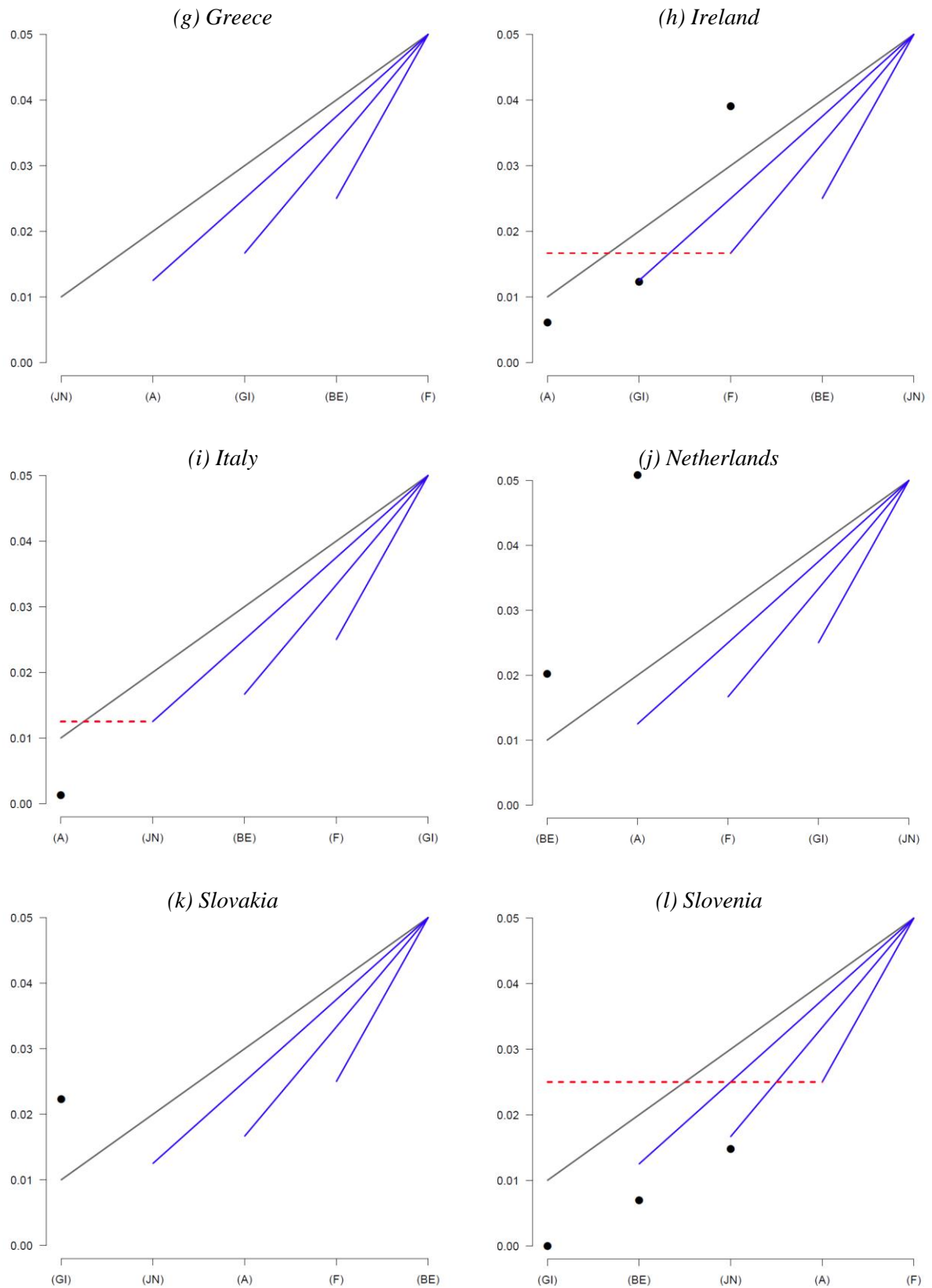


Fig. 1. Hommel's procedure by combining p -values across sectors (at the 5% significance level) (*continued*)

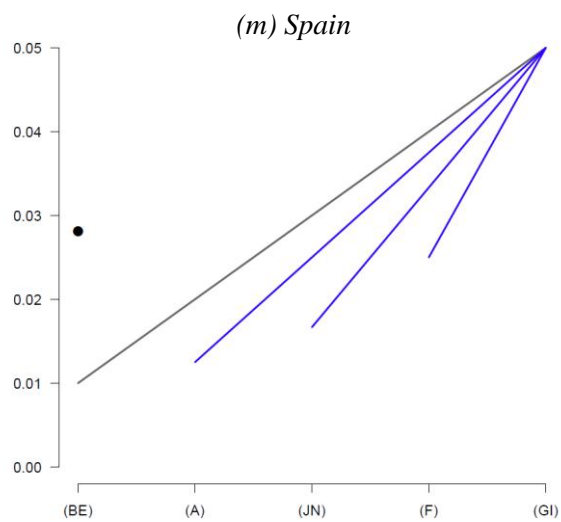


Fig. 1. Hommel's procedure by combining p -values across sectors (at the 5% significance level) (*continued*)

Notes: (A) - Agriculture, (BE) - Manufacturing, (F) - Construction, (GI) - Services, (JN) - Other Business Services.

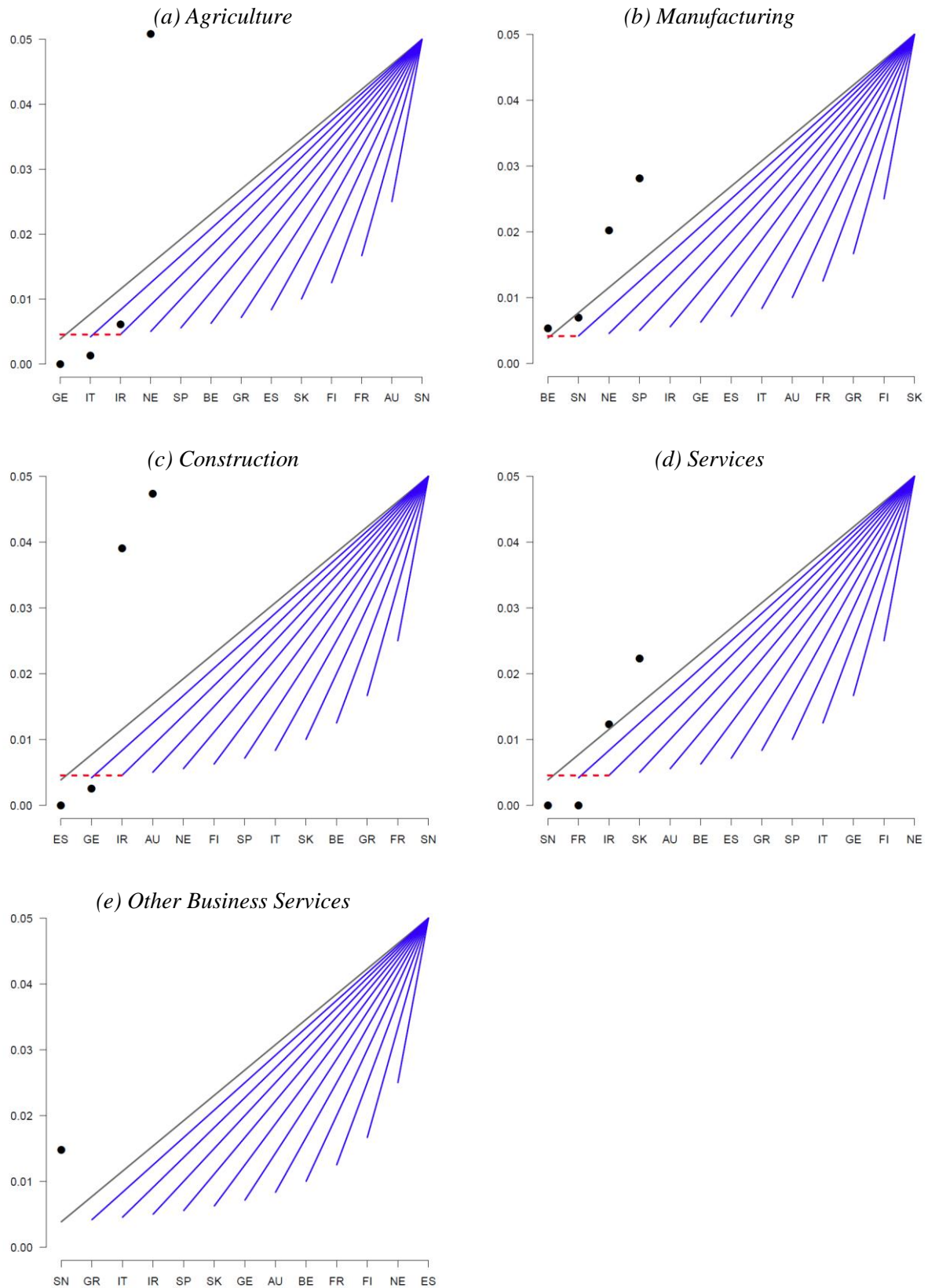


Fig. 2. Hommel's procedure by combining p -values across countries (at the 5% significance level)

Notes: BE – Belgium, ES – Estonia, FI – Finland, FR – France, GE – Germany, GR – Greece, IR – Ireland, IT – Italy, NE – Netherlands, SK – Slovakia, SN – Slovenia, SP – Spain.

Appendix A

Table A.1

Description of sectors used in estimation

	<i>Sector</i>	<i>Description</i>	<i>NACE Rev.2.</i>
1	Agriculture	Agriculture, forestry and fishing.	Section A
2	Manufacturing	Manufacturing (except construction).	Sections B to E
3	Construction	Construction.	Section F
4	Services	Wholesale and retail trade, transport, accommodation and food service activities	Sections G to I
5	Other business services	Information and communication, financial and insurance activities; real estate activities; professional, scientific and technical activities; administrative and support service activities.	Sections J to N

Table A.2

Descriptive statistics for countries

	<i>Mean</i>	<i>Std. Dev</i>	<i>Minimum</i>	<i>Maximum</i>	<i>Mean</i>	<i>Std. Dev</i>	<i>Minimum</i>	<i>Maximum</i>
	Inflation $\pi_{ij,t}$				Labour share $S_{ij,t}$			
Austria	0.004	0.016	-0.094	0.125	0.441	0.049	0.093	0.812
Belgium	0.002	0.017	-0.092	0.145	0.449	0.039	0.073	0.748
Estonia	0.011	0.056	-0.194	0.313	0.466	0.064	0.252	0.978
Finland	0.003	0.030	-0.243	0.282	0.470	0.039	0.135	0.786
France	0.004	0.019	-0.103	0.134	0.508	0.026	0.181	0.703
Germany	0.000	0.030	-0.349	0.182	0.541	0.044	0.256	0.883
Greece	0.005	0.055	-0.372	0.270	0.305	0.048	0.067	0.585
Ireland	0.003	0.147	-0.771	0.546	0.421	0.056	0.121	1.087
Italy	0.004	0.017	-0.163	0.089	0.412	0.085	0.167	1.023
Netherlands	0.004	0.043	-0.301	0.390	0.497	0.058	0.202	0.749
Slovakia	0.009	0.063	-0.295	0.256	0.368	0.059	0.193	0.578
Slovenia	0.008	0.043	-0.167	0.200	0.482	0.038	0.132	0.723
Spain	0.007	0.079	-0.207	0.208	0.419	0.041	0.113	0.637

Table A.3

Descriptive statistics for sectors

	<i>Mean</i>	<i>Std. Dev</i>	<i>Minimum</i>	<i>Maximum</i>	<i>Mean</i>	<i>Std. Dev</i>	<i>Minimum</i>	<i>Maximum</i>
	Inflation $\pi_{ij,t}$				Labour share $S_{ij,t}$			
Agriculture	0.002	0.110	-0.771	0.546	0.245	0.064	0.067	1.023
Manufacturing	0.003	0.032	-0.173	0.219	0.528	0.049	0.226	0.748
Construction	0.008	0.040	-0.430	0.313	0.550	0.065	0.193	1.087
Services	0.006	0.030	-0.192	0.218	0.541	0.045	0.241	0.778
Other Business Services	0.006	0.025	-0.170	0.132	0.360	0.025	0.197	0.601

Table A.4

List of variables used to estimate the factors

	Variable	Transformation	Source
1	Compensation of Employees	growth rate, q-o-q	Eurostat
2	Harmonized Index of Consumer Prices (2000=100)	growth rate, q-o-q	Eurostat
3	Domestic Demand, constant prices	growth rate, q-o-q	Eurostat
4	Exports of Goods and Services, constant prices	growth rate, q-o-q	Eurostat
5	Government Consumption Deflator	growth rate, q-o-q	Eurostat
6	Government Consumption, constant prices	growth rate, q-o-q	Eurostat
7	Gross Domestic Product, constant prices	growth rate, q-o-q	Eurostat
8	Gross Investment Deflator	growth rate, q-o-q	Eurostat
9	Gross Investment, constant prices	growth rate, q-o-q	Eurostat
10	Government Expenditure (ratio to GDP)	ratio to GDP	ECB
11	Government Expenditure	growth rate, q-o-q	ECB
12	Government Revenue (ratio to GDP)	ratio to GDP	ECB
13	Government Primary Surplus (ratio to GDP)	ratio to GDP	ECB
14	Imports of Goods and Services, constant prices	growth rate, q-o-q	Eurostat
15	Industrial Production – Manufacturing, constant prices	growth rate, q-o-q	Eurostat
16	Monetary Aggregate M1	growth rate, q-o-q	ECB
17	Monetary Aggregate M2	growth rate, q-o-q	ECB
18	Monetary Aggregate M3	growth rate, q-o-q	ECB
19	Nominal Effective Exchange Rate	growth rate, q-o-q	Eurostat
20	Brent Crude Oil 1 Month Forward (fob, per barrel)	growth rate, q-o-q	ECB
21	Private Consumption Deflator	growth rate, q-o-q	Eurostat
22	Private Consumption, constant prices	growth rate, q-o-q	Eurostat
23	Real Effective Exchange Rate	growth rate, q-o-q	Eurostat
24	Index of Retail Trade (2005=100)	growth rate, q-o-q	Eurostat
25	Trade Balance (ratio to GDP)	ratio to GDP	Eurostat
26	Unit Labour Costs	growth rate, q-o-q	Eurostat
27	Unemployment Rate	percentage points	Eurostat

Table A.5

Number of factors determined by information criterion

	<i>Number of factors</i>	<i>IC₂</i>
Austria	2	-0.014
Belgium	5	-0.049
Estonia	1	-0.181
Finland	7	-0.007
France	1	-0.149
Germany	1	-0.068
Greece	7	-0.017
Ireland	1	-0.195
Italy	1	-0.136
Netherlands	1	-0.061
Slovakia	1	-0.028
Slovenia	8	-0.132
Spain	1	-0.228

Note: IC_2 stands for information criterion as in Bai and Ng (2002).

Table A.6

The BSQT results for sector-specific inflation and labour shares

	Proportion of stationary units	Number of rejections	Number of time Observations	Number of Cross-sections
<i>Inflation $\pi_{ij,t}$</i>				
<i>Agriculture</i>	1.00	13	47	13
<i>Manufacturing</i>	1.00	13	47	13
<i>Construction</i>	1.00	13	47	13
<i>Services</i>	1.00	13	47	13
<i>Other Business Services</i>	1.00	13	47	13
<i>Labour share $S_{ij,t}$</i>				
<i>Agriculture</i>	0.75	10	48	13
<i>Manufacturing</i>	0.00	0	48	13
<i>Construction</i>	0.25	3	48	13
<i>Services</i>	0.50	7	48	13
<i>Other Business Services</i>	0.25	3	48	13

Notes: the quantiles to be tested are 0, 0.25, 0.5 and 0.75. The block length is chosen equal to the smallest integer larger than $1.75 T^{1/3}$. The number of bootstrap replications is 4999. The lag length for the Augmented Dickey-Fuller test is selected using Schwarz Bayesian information criterion, whereas the maximum lag length is set to the largest integer smaller than $12 (T/100)^{1/4}$. The unions of rejections approach (see Harvey, 2012) is applied in order to deal with uncertainty about deterministic components. The overall significance level is 5%.

Table A.7

Correlations between inflation and labour shares

	<i>Agriculture</i>	<i>Manufacturing</i>	<i>Construction</i>	<i>Services</i>	<i>Other Business Services</i>
Austria	-0.082	-0.069	-0.553	-0.163	-0.022
Belgium	-0.184	-0.131	-0.484	-0.305	-0.119
Estonia	-0.392	0.179	-0.480	-0.141	-0.231
Finland	-0.197	-0.010	-0.373	-0.009	-0.201
France	-0.091	-0.014	-0.381	-0.500	-0.169
Germany	-0.278	-0.095	-0.213	-0.101	-0.128
Greece	-0.041	-0.101	-0.141	-0.025	0.007
Ireland	-0.089	-0.008	-0.366	-0.311	-0.236
Italy	-0.185	-0.143	-0.251	-0.186	-0.409
Netherlands	-0.326	-0.170	0.304	-0.081	-0.118
Slovakia	0.178	0.018	-0.039	0.042	-0.143
Slovenia	-0.208	-0.319	-0.319	-0.451	-0.224
Spain	-0.302	-0.048	0.334	-0.130	-0.180
Spain	-0.082	-0.069	-0.553	-0.163	-0.022

Table A.8Results from Fisher-type test by combining p -values across sectors

	P_F	$\tau_{0.01}^*$	$\tau_{0.05}^*$	$\tau_{0.1}^*$	$p^*(P_F)$
Austria	5.334	16.025	11.404	9.228	0.344
Belgium	18.700**	21.453	16.599	14.259	0.026
Estonia	26.599***	25.430	19.115	16.402	0.007
Finland	4.368	16.739	11.565	9.316	0.466
France	10.695**	13.580	9.536	7.928	0.032
Germany	4.665	24.364	18.231	15.421	0.875
Greece	6.515	23.238	17.803	15.250	0.694
Ireland	12.543**	16.032	11.475	9.332	0.037
Italy	4.418	15.103	10.751	8.798	0.438
Netherlands	7.345	28.055	21.986	19.225	0.803
Slovakia	8.212	27.333	20.529	17.773	0.621
Slovenia	23.308**	25.371	19.187	16.441	0.017
Spain	19.927**	22.329	14.980	12.246	0.016

Notes: P_F denotes the Fisher-type test statistic, $\tau_{0.01}^*$, $\tau_{0.05}^*$ and $\tau_{0.1}^*$ are simulated critical values at the 1%, 5% and 10% significance levels respectively, $p^*(P_F)$ denotes the simulated p -values. In order to carry out the simulations to obtain the critical values for Fisher-type test, the cases for which the solution to the difference equation is indeterminate, *i.e.*, when $|\lambda_{ij}^f + \lambda_{ij}^b| > 1$ have been dropped out. *** and ** indicate statistical significance at the 1% and 5% levels respectively.

Table A.9Results from Fisher-type test by combining p -values across countries

	P_F	$\tau_{0.01}^*$	$\tau_{0.05}^*$	$\tau_{0.1}^*$	$p^*(P_F)$
<i>Agriculture</i>	8.438	28.917	22.219	19.608	0.745
<i>Manufacturing</i>	45.153***	40.869	33.392	29.931	0.005
<i>Construction</i>	36.199*	46.703	39.117	35.207	0.084
<i>Services</i>	41.893**	46.127	38.220	34.278	0.025
<i>Other Business Services</i>	20.946	37.671	30.129	26.455	0.273

Notes: P_F denotes the Fisher-type test statistic, $\tau_{0.01}^*$, $\tau_{0.05}^*$ and $\tau_{0.1}^*$ are simulated critical values at the 1%, 5% and 10% significance levels respectively, $p^*(P_F)$ denotes the simulated p -values. In order to carry out the simulations to obtain the critical values for Fisher-type test, the cases for which the solution to the difference equation is indeterminate, *i.e.*, when $|\lambda_{ij}^f + \lambda_{ij}^b| > 1$ have been dropped out. *** and ** indicate statistical significance at the 1% and 5% levels respectively.

Appendix B

The data is corrected for additive outliers employing a procedure based on wavelet analysis. The principal idea is to decompose the time series into different frequency components, which is in essence a band pass filtering¹².

We transform the data using a maximal overlap discrete wavelet transform (MODWT). The first level wavelet coefficients correspond to the vector of detail coefficients $\mathbf{D}_1 = (d_{1,1}, d_{1,2}, \dots, d_{1,N})$ and the approximation coefficients $\mathbf{A}_1 = (a_{1,1}, a_{1,2}, \dots, a_{1,N})$, where N is the number of wavelet coefficients. We use the Haar wavelet filter as there is no need to employ a smoother wavelet filter for outlier detection (see Grané and Veiga, 2010).

The first level detail coefficients \mathbf{D}_1 are very sensitive to any kind of abrupt changes in data, such as presence of outliers. Therefore, it is sufficient to screen only the finest-scale detail coefficients (see Bilen and Huzurbazar, 2002; Grané and Veiga, 2010). Following Bilen and Huzurbazar (2002), the detail coefficients in \mathbf{D}_1 are compared with a threshold limit given by

$$\tau_1 = \hat{\sigma}_1 \sqrt{2 \ln(N)}, \quad (\text{B.1})$$

where $\hat{\sigma}_1$ is the mean of the absolute deviations from the median based on the finest-scale detail coefficients

$$\hat{\sigma}_1 = N^{-1} \sum_{n=1}^N |d_{1,n} - M_1|, \quad (\text{B.2})$$

where M_1 is the median of \mathbf{D}_1 . Outlier correction is carried out by setting to zero those detail coefficients which exceed the threshold limit τ_1 . Formally, a hard-thresholding procedure is performed and the modified finest-scale detail coefficients $\tilde{\mathbf{D}}_1 = (\tilde{d}_{1,1}, \tilde{d}_{1,2}, \dots, \tilde{d}_{1,N})$ are obtained as $\tilde{d}_{1,n} = 0$ if $|d_{1,n}| > \tau_1$ and $\tilde{d}_{1,n} = d_{1,n}$ otherwise for $n = 1, \dots, N$. Finally, the time series are reconstructed by applying inverse maximal overlap discrete wavelet transform (IMODWT) using the modified detail coefficients $\tilde{\mathbf{D}}_1$ and the original approximation coefficients \mathbf{A}_1 .

¹² For an extensive discussion about wavelet analysis we refer to Gençay *et al.* (2002).

Appendix C

Under the null hypothesis that $\varphi_{ij} = 0$, the infinite sum of expected real marginal costs pre-multiplied by φ_{ij} is zero. The rational expectations solution in Eq. (4) will thus reduce to

$$\pi_{ij,t} = \rho_{ij,1}\pi_{ij,t-1} + (1 - \rho_{ij,1}\lambda_{ij}^f)^{-1}\varepsilon_{ij,t}. \quad (\text{C.1})$$

Hence, if the null hypothesis is true, the future real marginal costs do not affect inflation and inflation dynamics are characterised by an autoregressive process of order one.

We test the global null hypothesis that $\varphi_{ij} = 0$ for all j given i (or $\varphi_{ij} = 0$ for all i given j) using a Fisher-type test (see Fisher, 1932). The Fisher-type test statistic is

$$P_F = -2 \sum_{n=1}^N \ln(p_n), \quad (\text{C.2})$$

where p_n stands for p -value associated with computed t statistic of individual test n , and N is the number of individual tests. The Fisher-type test is only valid, however, under the assumption that the individual tests are independent. If the cross-sectional units are dependent the Fisher-type test statistic does not follow χ^2 distribution under the null hypothesis, and the critical values of the test therefore have to be bootstrapped.

The algorithm used to obtain the critical values for Fisher-type test is as follows.

- i. Using the estimated parameters of the hybrid New Keynesian Phillips curve in Eq. (3) we solve for the characteristic roots $\hat{\rho}_{ij,1}$ and $\hat{\rho}_{ij,2}$ of the difference equation for inflation.
- ii. Bootstrap errors $\hat{\varepsilon}_{ij,t}^*$ are generated using the estimated residuals $\hat{\varepsilon}_{ij,t}$ by drawing overlapping blocks at random with replacement. In each bootstrap replication, the same randomly generated index is used for each country i and sector j so that dependence is preserved across cross-sectional units. The block length is set equal to 4.
- iii. The bootstrap sample of inflation is generated by $\pi_{ij,t}^* = \hat{\rho}_{ij,1}\pi_{ij,t-1}^* + (1 - \hat{\rho}_{ij,1}\hat{\lambda}_{ij}^f)^{-1}\hat{\varepsilon}_{ij,t}^*$ with $\pi_{ij,0}^* = 0$. The first 100 simulated values are discarded and the bootstrap sample size T is set equal to the actual sample size.
- iv. The time series for the real marginal cost is assumed to be generated by an autoregressive process of order one¹³. We then fit the $AR(1)$ model $s_{ij,t} = \varrho_{ij}s_{ij,t-1} + \eta_{ij,t}$ and obtain parameter estimate $\hat{\varrho}_{ij}$ and sample variance of

¹³ This assumption is common in the literature, see for example Byrne *et al.* (2013).

estimated residuals $\hat{\sigma}_{ij}^2 = \text{var}(\hat{\eta}_{ij,t})$. The simulated sample for $s_{ij,t}^*$ is then generated as $s_{ij,t}^* = \hat{q}_{ij} s_{ij,t-1}^* + \hat{\eta}_{ij,t}^*$, where the errors $\hat{\eta}_{ij,t}^*$ are drawn from the normal distribution $N(0, \hat{\sigma}_{ij}^2)$.

- v. In each bootstrap replication for each country i and sector j , the sectoral hybrid NKPC is estimated using the simulated data series $\pi_{ij,t}^*, s_{ij,t}^*$. In line with Inoue and Shintani (2006), the lags (from the 2nd to 8th) of simulated inflation are used as instruments in the bootstrap for GMM estimation.
- vi. In total, $B = 9999$ simulation replications are run. In each replication, b , the simulated test statistic P_F^{*b} is obtained based on p -values corresponding to t statistic of individual significance test on φ_{ij} .
- vii. The simulated test statistics are then ordered from smallest to largest and the critical value τ_α^* is defined as the $1 - \alpha$ quantile for empirical distribution of the test statistic. The p -value based on the simulated test statistics is $p^*(P_F) = B^{-1} \sum_{b=1}^B \mathbb{I}(P_F^{*b} > P_F)$, where $\mathbb{I}(\cdot)$ is an indicator function equal to 1 if the argument is true and 0 otherwise.