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Non-linearities in euro area inflation persistence

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A B S T R A C T

This paper investigates the nature of inflation dynamics with a special focus on inflation persistence. Using data from euro area member-states we estimate dynamic non-linear panel models addressing in detail econometric issues concerning unobserved heterogeneity, genuine state dependence, and the initial conditions problem. After controlling for observed and unobserved heterogeneity, our results suggest that the degree of inflation persistence is genuine and varies depending on whether the inflation rate is too high, within the range of ECB's target of price stability, too low or negative. This implies that policies to stabilize inflation in the short run will have longer-run effects.

1. Introduction

Over the last four decades inflation dynamics in many European countries have undergone tremendous changes regarding average inflation levels, persistence and volatility. The ‘Great Inflation’ period of the 1970s and early 1980s was succeeded by a decade of declining inflation rates and progressively reduced volatility (middle 1980s to middle 1990s). The changes in inflation behaviour became particularly pronounced for most member-states of the European Monetary Union (EMU) soon after the introduction of the euro. The framework for the conduct of monetary policy along with the European Central Bank's (ECB) policy strategy proved quite successful in taming consumer-price inflation and anchoring inflationary expectations. However, low-inflation environments face the possibility of deflationary episodes as moderate fluctuations around a low level of inflation can turn inflation to deflation (Bordo and Filardo, 2005). This is exactly what happened to a certain number of euro area countries after the outbreak of the 2008 economic crisis.

Against this background it is of great interest to ask whether deflation episodes can contribute to further downward price pressures turning deflation itself into a more permanent situation. This amounts to asking whether negative inflation exhibits the same pattern of persistence as positive inflation does. Though resolving the uncertainty surrounding this issue is of paramount importance for the conduct of monetary policy, the literature has not provided a definite answer. However, certain theoretical arguments have been proposed in justifying the non-linearities of inflation persistence. Low competition in product markets, as well as rigid labour markets, allow firms to reset their prices upwards during good times and to delay a (downward) price adjustment during periods of economic slack. In such circumstances firms tend to be more responsive to negative supply shocks (e.g. firms set higher prices when confronted with higher input prices) and less responsive to positive supply shocks (e.g. prices may be left unchanged after a decrease in input prices), or to react to product and labour market slack only after economic activity measures (e.g. unemployment) have reached a certain threshold value (Barnes and Olivei, 2003; Peach et al., 2011).[1] This firms' behaviour possibly explains why inflation persists at disproportionally high levels during periods of low economic activity. Moreover, the services' sector, which prevails in EU countries, differs from the rest productive sectors of the economy in two important aspects: services are largely non-tradable and labour intensive. The first feature makes firms of the services sector immune to

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[1] The weakness of linear Phillips curves to describe the linkage between inflation and unemployment for certain time intervals during the “Great Moderation” era of the US economy led Barnes and Olivei (2003) and Peach et al. (2011) to adopt a piecewise linear specification of the Phillips relationship. Using US data, the two papers show that inflation responds asymmetrically to unemployment changes depending on the level of unemployment.

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international price competition and thus less prone to price declines. The second feature adds to downward price rigidity through downward wage rigidity.[2]

Motivated by the idea that inflation persistence might not be symmetric across different states of inflation, this paper investigates the existence of non–linearities in the responsiveness of current inflation to its own lag for euro area countries. From this perspective, it belongs to a class of papers that search for various forms of instability in inflation dynamics using as building blogs traditional Phillips curve relationships (e.g. Laxton et al., 1999; Aguiar and Martins, 2005; Baghi et al., 2007; Musso et al., 2009). So, while we are not the first to address the issue of non-linearity, this is the first attempt that explicitly links variations in inflation inertia with different ranges of the inflation level, namely, deflation, too-low inflation, price stability, and medium to high inflation.

Another contribution of the paper lies in the adopted methodological approach. Specifically, we employ a dynamic random effects ordered probit framework, which allows us to capture the presence of asymmetric features in the response of current inflation with respect to its own lag. Our results are consistent with the idea that the relationship of current inflation to its own lag varies, depending on whether the inflation rate is too high, within the range of ECB's target of price stability, too low or negative.

The paper proceeds as follows. Section 2 presents the data used and the empirical framework, Section 3 considers the empirical results and their implications and investigates the robustness of our results to alternative model specifications, while Section 4 concludes.

2. Data and methodology

Our country sample includes the 11 European countries that have been full members of EMU since 1999.[3] All data used in the estimations is obtained from Eurostat and is of quarterly frequency, spanning the period 1997:Q1 to 2015:Q3. The measure of inflation is the annualized Harmonized Index of Consumer Prices (HICP hereafter) inflation rate. We prefer headline over core inflation because the inflation target of the ECB is explicitly stated in terms of headline measures and its policy makers pay less attention to core measures. Economic slack is proxied by the seasonally adjusted unemployment rate to eliminate measurement problems and uncertainty surrounding alternative proxies like the output gap.

In the empirical literature of inflation persistence there are two main methodological approaches as to measure persistence. The first and most common methodology utilizes a simple univariate time-series framework and assumes that inflation follows an autoregressive process of order $p(\mathrm{AR}(p))$. From this model various measures of inflation persistence, such as the "sum of autoregressive coefficients", the "spectrum at zero frequency", the "largest autoregressive root" and the "half-life", can be derived. The second approach utilizes multivariate econometric models and assumes that inflation depends not only on its own lag but on other variables as well. The advantage of the multivariate approach is that it offers a deeper analysis of persistence, since it incorporates other economic variables that affect the evolution of inflation. In this paper we use the multivariate approach and in particular a dynamic Phillips curve framework. Moreover, since we are examining a set of countries under a single central bank rather than a single country, time-series analysis is not appropriate and we use longitudinal models, which among others eliminate country heterogeneity.

In its simplest form the dynamic Phillips curve assumes that the current level of inflation ($HICP_{it}$) depends on its own lag ($HICP_{it-1}$) and the current level of unemployment (u_{it}) as well as other explanatory variables (χ_t). This model for country $i=1,\ldots,N$ in time $t=2,\ldots,T$ takes the form of

$$HICP_{it} = cHICP_{it-1} + \theta u_{it} + \chi_t'\beta + \alpha_i + \varepsilon_{it}. \tag{1}$$

In Eq. (1) the level of inflation persistence is proxied by the size of coefficient c, which by construction is assumed to be constant. Our main argument in this paper is that inflation persistence is not constant, i.e. not linear, but varies depending on the level of previous quarter's inflation. In order to test our assumption we need to distinguish among different levels of inflation, namely disinflation, low inflation, inflation around the target set by ECB and high inflation. One way to do this is to construct an ordered variable representing the four aforementioned levels of inflation for each country i in period t and use this to estimate a dynamic Phillips curve instead of Eq. (1). Our new dependent variable is:

$$\pi_{it} = \begin{cases} 1 \text{ if } HICP_{it}<0 \\ 2 \text{ if } 0\le HICP_{it}<1 \\ 3 \text{ if } 1\le HICP_{it}<2.5 \\ 4 \text{ if } 2.5\le HICP_{it} \end{cases} \tag{2}$$

Identifying true inflation persistence, i.e. the effect of previous inflation status on the probability of current inflation status, as opposed to heterogeneity, suggests a modelling approach that incorporates both observable and unobservable influence on inflation. Since the level of inflation is an ordered variable, the dynamic random effects ordered probit framework represented by equation below is the most appropriate.[4]

$$\pi_{it}^* = \gamma_1\pi_{it-1}^1 + \gamma_2\pi_{it-1}^2 + \gamma_4\pi_{it-1}^4 + \theta u_{it} + \chi_t'\beta + \alpha_i + \varepsilon_{it} \tag{3}$$

The subscript $i = 1,\ldots,N$ denotes countries that are included in our sample and the subscript $t = 2,\ldots,T$ represents the time periods for which the model is estimated. π_{it} is an ordinal variable representing the level of inflation and takes the values {1, 2, 3, 4} depending on the value of π_{it}^*, a latent measure of the level of inflation accordingly to Eq. (2). u_{it} is the level of unemployment and x_t contains strictly exogenous variables. In particular it includes year dummies to capture any trend effect, as well as an indicator variable of whether the country has physically adopted Euro as its currency. Obviously $\pi_{it-1}^{j=1,2,4}$ is the level of inflation of country i in the previous quarter. The random error term in this model is composed of two terms. The country specific error term α_i captures unobserved heterogeneity which differs between countries but remains constant for each country,[5] while ε_{it} is the usual error term with zero mean, uncorrelated with itself, with x_{it} and α_i as well as homoscedastic.

In these models special attention should be paid to the treatment of the initial conditions problem, which arises when the beginning of the examined period does not coincide with beginning of the stochastic data generating process. More specifically in a dynamic random effects ordered probit model, the presence of the lagged dependent variable means that there is a correlation induced between the first observation of dependent variable π_{i1} and the unobserved heterogeneity α_i. To treat the initial conditions issue we adopt the solution suggested by Wooldridge (2005).[6] Wooldridge suggests using a conditional

[2] A number of papers explain inflation's persistence through indexation of price contracts (Christiano et al., 2005), rule-of thumb behaviour (Gali and Gertler, 1999) or alternative contract assumptions (Fuhrer and Moore, 1995). See, Fuhrer (2011) and Woodford (2007) for a review of the related literature.

[3] These are Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, the Netherlands, Portugal and Spain.

[4] The choice of an ordered model stems from the nature of our dependent variable while the choice of random effects comes from the fact that in non-linear models fixed effects are problematic. Maximum likelihood estimator is inconsistent in probit models with fixed effects because it suffers from incidental parameter problem (Neyman and Scott, 1948).

[5] In the random-effects models, it is assumed that α_i in Eq. (3) is purely random. This assumption implies that α_i is uncorrelated with the regressors.

[6] In the literature there are two other solutions proposed by Heckman (1981a, 1981b) and Orme (1996) to the problem of initial conditions. Both involve a separate equation for the initial period and need proper instrument(s) for identification, which should determine initial period's inflation but not subsequent. As such instrument is difficult to find we apply Wooldridge's estimator. Arulampalam and Stewart (2009) show that all three estimators provide similar results and none consistently performs better than the others.

Table 1
Inflation and unemployment descriptive statistics, 1997:Q1 to 2015:Q3.

Country	Unemployment rate (period average) [1]	Inflation rate (period average) [2]	Inflation regimes (frequencies) *HICP* < 0	0 ≤ HICP < 1	1 ≤ *HICP* < 2.5	2.5 ≤ *HICP*	Correlation between [1] & [2]
Austria	4.85 (0.59)	1.78 (0.87)	1	10	51	13	−0.257*
Belgium	8.02 (0.78)	1.88 (1.16)	5	9	40	21	−0.537***
Finland	8.81 (1.47)	1.80 (1.06)	3	12	39	21	−0.369**
France	9.29 (0.99)	1.55 (0.86)	3	17	49	6	−0.667***
Germany	7.92 (1.89)	1.47 (0.79)	2	17	50	6	0.061
Ireland	8.10 (4.02)	2.01 (1.92)	8	12	25	30	−0.748***
Italy	9.29 (1.98)	2.04 (0.94)	2	9	42	22	−0.462***
Luxembourg	4.13 (1.32)	2.28 (1.37)	5	7	29	34	−0.067
Netherlands	5.10 (1.26)	2.01 (1.20)	2	9	43	21	−0.536***
Portugal	9.46 (3.53)	2.17 (1.41)	9	8	24	34	−0.493***
Spain	15.64 (5.90)	2.35 (1.41)	7	5	24	39	−0.621***
Total	8.24(3.99)	1.94 (1.25)	47	115	416	247	−0.269***

Notes: Standard deviation in parentheses.
* p-Value <0.10.
** p-Value <0.05.
*** p-Value <0.01.

maximum likelihood estimator and modelling the distribution of inflation from the second until the final quarter conditioning on the explanatory variables and the inflation status in the initial year. According to Wooldridge unobserved heterogeneity can be modelled conditional on initial inflation π_{i1}. Thus, Eq. (3), becomes:

$$\pi^*_{it} = \gamma_1\pi^1_{it-1} + \gamma_2\pi^2_{it-1} + \gamma_4\pi^4_{it-1} + \theta u_{it} + \chi'_t\beta + \delta_1\pi^1_{i1} + \delta_2\pi^2_{i1} + \delta_4\pi^4_{i1} + \alpha_i + \varepsilon_{it}. \quad (4)$$

This is a simple conditional dynamic random effects ordered probit, easily estimated, with the initial inflation as an extra set of regressors. A test of exogeneity of the initial conditions can be performed by a simple t-test on the coefficients of the initial inflation $\pi^{j=1,2,4}_{it-1}$.

Since we argue that the level of state dependence in inflation is not the same, but depends on both the origin as well as the destination, we also estimate transition probabilities between the different inflation regimes. These conditional probabilities are the sample average of the probability of belonging to inflation regime j given that the previous inflation regime was m, while keeping the rest of the regressors in their observed values. Moreover, one needs to appropriate scale predicted probabilities taking into consideration that these stem from a random effects model (Arulampalam, 1999; Wooldridge, 2005). Transition probabilities are defined accordingly to Eq. (5) below:

$$\hat{P}(\pi_{it} = j|\pi_{it-1} = m, U = u, X = x) = \frac{1}{N}\sum_{i=1}^{N}\Phi\left\{\frac{\left(\mu_j - \hat{\gamma}_m - \hat{\theta}u_{it} - \chi'_t\hat{\beta}\right)}{\sqrt{(1+\sigma^2_a)}}\right\} - \Phi\left\{\frac{\left(\mu_{j-1} - \hat{\gamma}_m - \hat{\theta}u_{it} - \chi'_t\hat{\beta}\right)}{\sqrt{(1+\sigma^2_a)}}\right\} \quad (5)$$

where $\Phi(\cdot)$ is the standard normal distribution function and μ_j are the cut points as estimated in Eq. (4). Since we cannot separately identify an intercept and the cut points we have adopted a conventional normalization and set the constant term equal to zero.[7]

Our dynamic model also allows us to distinguish between aggregate state dependence (ASD) and genuine state dependence (GSD). ASD is the simple conditional probability of remaining two consecutive quarters in the same inflation regime, as expressed in Eq. (5). However, this might be an effect of observed and unobserved heterogeneity (Heckman, 1981a, 1981b). Genuine state dependence arises when the chances of being in a particular inflation regime this period depend on being in the same inflation regime in the previous period, controlling for country heterogeneity (observed and unobserved). For example, the experience of high inflation might be the result of fiscal policies through taxation, lowering the chances that a country with given attributes escapes high inflation in the future. One method to distinguish between spurious and GSD is to look at the change in the predicted probabilities conditional on previous inflation status. This is the standard marginal effect calculation that also accounts for the distribution of unobserved heterogeneity in the sample. For every country, predicted probabilities for each inflation regime are calculated conditional, first on being in that regime the previous quarter and secondly on being in the regime of stable inflation ($\pi_{it-1} = 3$). The difference between the first and the second, averaged over the sample, gives an estimate of GSD or equivalently the persistence effect:

$$GSD_{\pi=j} = \hat{P}(\pi_{it} = j|\pi_{it-1} = j, U = u, X = x) - \hat{P}(\pi_{it} = j|\pi_{it-1} = 3, U = u, X = x). \quad (6)$$

This is actually the probability of observing a randomly chosen country in a particular inflation regime in the current period conditional on previous regime of stable inflation. It is clear that in our ordered model with four categories the estimation of marginal effects is estimated on the basis of a reference category, in this case $\pi_{it\,-} = 3$, which implies that our measure of genuine state dependence can only be expressed for $j \neq 3$ inflation regimes.

3. Results

Table 1 provides some insight to the data used in the analysis and their interrelation.[8] While the inflation rate during the examined period was on average around 2% and the unemployment rate around 8%, significant cross country variations exist. The mean unemployment rate records its lowest value in Luxembourg (4.13%) and its highest in Spain (15.64%). On average the inflation rate ranges between 1.47% in Germany and 2.35% in Spain. Even though the majority of the examined countries record inflation rate between 1 and 2.5% there is a significant number of cases where the inflation rate is higher than 2.5. Moreover, for 47 cases the inflation rate is below zero and for 115 it is close to zero. Moreover, it is worth mentioning that, with the exception of Germany and Luxembourg, the correlation between inflation and unemployment is negative and highly significant.[9] This for some countries is quite strong (Ireland: −0.748) and for some relatively weak (Austria:

[7] All models were estimated using Stata's *oprobit* and *xtoprobit* build in commands (StataCorp, 2013).

[8] A graphical presentation of the evolution and interrelation of inflation and unemployment is presented in the Appendix A.

[9] The correlation between inflation and unemployment is negative for Luxembourg and positive for Germany but in both not statistically different from zero.

Table 2
Dynamic ordered probit of inflation probability.

	(1)	(2)	(3)
	Pooled model	Random effects	Random effects with IC
	NT = 814	NT = 814	NT = 814
Previous inflation			
$\pi < 0$	-2.169^{***}	-2.021^{***}	-2.094^{***}
	(0.258)	(0.257)	(0.260)
$0 \le \pi < 1$	-1.652^{***}	-1.505^{***}	-1.548^{***}
	(0.177)	(0.177)	(0.179)
$2.5 \le \pi$	1.858^{***}	1.638^{***}	1.659^{***}
	(0.139)	(0.149)	(0.148)
Unemployment rate	-0.0368^{***}	-0.0816^{***}	-0.0911^{***}
	(0.0129)	(0.0267)	(0.0240)
Adopted euro	-1.214^{***}	-1.252^{***}	-1.280^{***}
	(0.356)	(0.353)	(0.358)
Initial period's inflation			
$0 \le \pi < 1$			0.152
			(0.248)
$2.5 \le \pi$			0.708^{***}
			(0.241)
Cut 1	-3.559^{***}	-3.895^{***}	-3.932^{***}
	(0.336)	(0.419)	(0.384)
Cut 2	-1.651^{***}	-2.050^{***}	-2.038^{***}
	(0.280)	(0.372)	(0.332)
Cut 3	1.414^{***}	0.965^{***}	1.075^{***}
	(0.288)	(0.358)	(0.324)
Log likelihood	−466.6	−464.2	−459.4
AIC	983.263	980.388	974.896
Wald test of non-linear inflation persistence			
$H_0: \gamma_1 = \gamma_2 = \gamma_4$			
χ^2	268.53	197.33	205.49
p-Value	[0.000]	[0.000]	[0.000]

Notes: (i) The table reports coefficients from a series of dynamic ordered models; model in column 3 allows for endogenous initial conditions using a methodology due to Wooldridge (2005). For estimation methods see text. (ii) Standard errors are in parentheses. (iii) All models also contain year dummies (not reported for brevity). (iv) Log likelihood and sample sizes refer to period 2 to T. (v) Cut 1–3 are the estimated cut points. (vi) AIC refers to Akaike's information criterion = −2 ∗ Log likelihood + 2 ∗ parameters.
*** *p*-Value <0.01.

−0.257). The differences in the inflation rate as well as the observed cross country variations provide motivation to examine the level and process of inflation mobility.

Table 2 presents the coefficient estimates for the ordered probit models based on pooled and random effects specifications. Random effects models were estimated using adaptive Gauss-Hermite quadrature. Since different scaling of the error variance is employed in the different models the estimated coefficients for the random effects model are properly scaled to be comparable to those reported for the pooled model. As a baseline we estimated a dynamic pooled model, column (1). Models in columns (2) and (3) introduce explicit unobserved heterogeneity into the dynamic model by specifying random effects and in addition model in column (3) models the initial conditions following Wooldridge (2005). Allowing for heterogeneity and taking into consideration initial conditions, i.e. the initial observed level of inflation of each country, improves the fit of the model as evidenced by the change in log-likelihood and the Akaike's information criterion (AIC).

To formally test our hypothesis that the level of inflation persistence depends on the level of previous inflation we estimated dynamic models which included lags of the categories of the dependent variable. In all specifications these are statistically significant. Moreover, for all three models a Wald test is estimated testing the null hypothesis that the coefficients of the lagged dependent variables are equal. In all cases the *p*-value of the Wald test strongly rejects the null hypothesis suggesting that the level of inflation persistence is not constant, but depends on the level of inflation the previous quarter. In addition, the last model uses a vector of dummy variables to represent the first-period observations on the dependent variable in order to model the initial conditions as suggested by Wooldridge (2005). Even though not all of them are significant their joint significance is very high. It is worth noting that there is a positive gradient in the estimated effects as we move from low to higher inflation. This implies that there exists a positive correlation between the initial period observations and unobserved latent inflation. Moreover, the relationship between inflation and unemployment in all specifications is, as expected, negative and highly significant and its effect is increasing as we control for unobserved heterogeneity and the initial conditions. Also an interesting finding is that the adoption of euro decreases the probability of higher inflation. This probably reflects ECB's effectiveness in taming inflation.

Table 3
Dynamic linear model of inflation.

	(1)	(2)	(3)
	OLS	Fixed-effects regression	Arellano-Bond dynamic panel-data estimation
	NT = 814	NT = 814	NT = 803
Previous inflation	0.787^{***}	0.729^{***}	0.723^{***}
	(0.0334)	(0.0376)	(0.0205)
Unemployment rate	−0.0116	-0.0442^{***}	-0.0490^{***}
	(0.00686)	(0.00809)	(0.00787)
Adopted euro	-0.240^{**}	-0.305^{**}	-0.285^{**}
	(0.0979)	(0.109)	(0.114)
Constant term	0.379^{***}	0.773^{***}	0.800^{***}
	(0.0857)	(0.0662)	(0.121)
Log-likelihood	−577.9	−561.8	
AIC	1197.79	1165.58	

Notes: (i) The table reports coefficients from a series of dynamic linear models estimating Eq. (1). (ii) Standard errors are in parentheses. (iii) All models also contain year dummies (not reported for brevity). (iv) Log likelihood and sample sizes refer to period 2 to T. (v) AIC refers to Akaike's information criterion = −2 ∗ Log likelihood + 2 ∗ parameters.
** *p*-Value <0.05.
*** *p*-Value <0.01.

Another test to examine whether there are significant inflation nonlinearities is to estimate a dynamic model on the observed level of current inflation rate similar to that presented in Eq. (1) and compare it with the ordinal models. Here the dependent variable is the observed HICP and it depends on its observed level in the previous quarter as well as on the other variables used in the ordered models (unemployment rate, adoption of euro and time). In this model specification inflation persistence is also approximated by the coefficient of the lagged dependent variable *c*. Since the dependent variable is now a continuous variable and in order for the results to be comparable with those in Table 2 we estimate Eq. (1) using (i) a pooled ordinary least squares (OLS), (ii) a fixed effects panel regression and (iii) Arellano-Bond's dynamic panel-data estimation and the results for these three dynamic linear models are presented in Table 3.[10] In terms of model choice using both the log-likelihood and the AIC, the non-linear models perform much better and have a better fit, providing an additional argument than inflation persistence is non-linear.

All coefficients in all three models have the expected sign and are statistically significant. The level of inflation persistence as it is measured by the coefficient of the lagged dependent variable is quite high and definitely higher than that from the non-linear models of Table 2. However, one should keep in mind that the non-linear models examine the probability of belonging in an inflation regime while the linear models examine the level of inflation per se. In the pooled OLS the level of persistence is

[10] In the estimation of the ordered models we use random effects due to computational reasons, however, in the linear models we use fixed effects as they are consistent. Moreover, a Hausman test suggests to prefer fixed over random effects.

Table 4
Inflation persistence, and tests of equality of inflation persistence between models.

	$P(\pi_t=1 \mid \pi_{t-1}=1) - P(\pi_t=1 \mid \pi_{t-1}=3)$		$P(\pi_t=2 \mid \pi_{t-1}=2) - P(\pi_t=1 \mid \pi_{t-1}=3)$		$P(\pi_t=4 \mid \pi_{t-1}=4) - P(\pi_t=1 \mid \pi_{t-1}=3)$	
Ordered models						
[1] Pooled ordered probit	15.08		29.84		50.58	
[2] Panel ordered probit	13.98		33.82		44.06	
[3] Panel ordered probit (Wooldridge, 2005)	14.06		25.03		43.64	
Linear models						
[4] OLS			78.72			
[5] FE			72.85			
[6] Arellano-Bond dynamic estimator			72.25			
Test	x^2	*p*-Value	x^2	*p*-Value	x^2	*p*-Value
H_0: [1] = [4]	311.25	[0.000]	110.79	[0.000]	60.06	[0.000]
H_0: [2] = [5]	277.03	[0.000]	108.51	[0.000]	43.67	[0.000]
H_0: [3] = [6]	299.45	[0.000]	109.15	[0.000]	45.87	[0.000]

Notes: (i) The table reports measures of inflation persistence estimated from a series of dynamic ordered (Table 2) and linear models (Table 4). (ii) All measures of inflation persistence are statistically significant at 1%.

0.787, quite higher than that of fixed effects, 0.729. This suggests that accounting for country heterogeneity is very important. Using the Arellano-Bond estimator, which takes initial conditions into consideration, the effect of previous inflation is even smaller 0.723.

To formally test that inflation persistence is non-linear, we perform a Wald test that the marginal effects from the ordered models are equal to coefficient of the corresponding linear model. The estimated inflation persistence is shown in Table 4, together with test statistics for the equality of linear and ordinal models. In all specifications estimated inflation persistence is positive and statistically significant; indicating that current inflation level or regime has an effect on the level of future inflation level or regime. In the linear models inflation persistence is found higher than in ordinal models.

Overall, the data support the proposed estimation strategy, i.e. accounting for different level of inflation persistence depending on the magnitude of previous inflation. According to the tests reported at the bottom of Table 4, in none of the three models that assume non-linear inflation persistence the level of persistence is equal to that of the corresponding model that assumes linear persistence: the null hypothesis that the marginal effects of the ordinal models are equal to the coefficients of the linear models is rejected at conventional levels.

To provide a better understanding of the association between current and previous level of inflation Table 5 presents the estimated transition probabilities of inflation as defined in Eq. (5) in the form of a transition matrix. The elements of the main diagonal correspond to the probabilities of recording the same state of inflation (ASD); those above the diagonal correspond to the probabilities of reporting a higher inflation level; while those below the diagonal correspond to the probabilities of reporting a lower inflation level.

Table 5
Inflation transition probabilities.

Initial ($t-1$) regime	Destination (t) regime probabilities: Inflation < 0	0 ≤ inflation < 1	1 ≤ inflation < 2.5	2.5 ≤ inflation
Inflation < 0	14.8%	45.6%	39.4%	0.3%†
0 ≤ inflation < 1	8.1%	36.3%	54.5%	1.0%
1 ≤ inflation < 2.5	0.7%	11.3%	73.5%	14.5%
2.5 ≤ inflation	0.0%†	1.1%	40.8%	58.1%

Notes: (i) The table reports conditional probabilities (Eq. (5)) from model in column 3 of Table 2. (ii) †Indicate insignificant transition probability at 5%.

As is evident from the results, inflation exhibits a tendency to stay close to its origin. However, the degree of persistence varies, depending on the inflation's own lag. For instance, when inflation lies close to the ECB's inflation target rate, that is around 2%, the probability that it will remain in this region is 73.5%. The probability of inflation persistence drops to 58.1% when prices are in the inflationary zone. Regarding the regimes of too-low and of negative inflation, the probabilities of inflation persistence take considerably lower values (36.3% and 14.8% respectively). In every case, the significant heterogeneity in the degree of inflation persistence across different inflation zones implies a short-run asymmetric response of inflation with respect to its own lag.

The estimated transition probabilities further demonstrate that large 'jumps' of inflation are highly unlikely, suggesting a smooth behaviour of inflation. Moreover, the estimated transition probabilities that lie bellow the main diagonal are typically lower than those lying above it, suggesting that inflation movements towards higher levels are more likely than movements towards lower ones. So, while deflation episodes or movements from the 'stable' inflation regime to disinflation cannot be ruled out, the estimated probabilities attached to such cases are very small (8.1% and 11.3% respectively). On balance, it seems that the higher the gap between the actual inflation rate and the ECB's target rate is, the higher the tendency of inflation to 'correct' itself and move towards the desired level is.

An intuitively appealing and plausible explanation for these results lies in ECB's credible monetary policy and its successful communication to economic agents of the euro area. Beyond the well-anchored inflation expectations to the ECB's target, our results are consistent with micro evidence provided by Inflation Persistence Network (IPN) of the ECB regarding firms price setting behaviour in the Euro area.[11] Specifically, according to IPN surveys, the price adjustment process of euro area firms takes place in two steps: the price review and the actual price change. Most firms review price one to three times per year and proceed to price resetting once a year. Combining the frequency of price reviews and actual price changes with the low frequency of quarterly deflation of our data sample suggests that firms will detect general price decreases rarely. Hence, firms may leave their prices unchanged after a detection of a deflation episode either because their previous price reviews revealed positive inflation or because they adopt a wait-and-see behaviour until their next price review.

[11] For a summary of the empirical evidence gathered by IPN, see Altissimo et al. (2006).

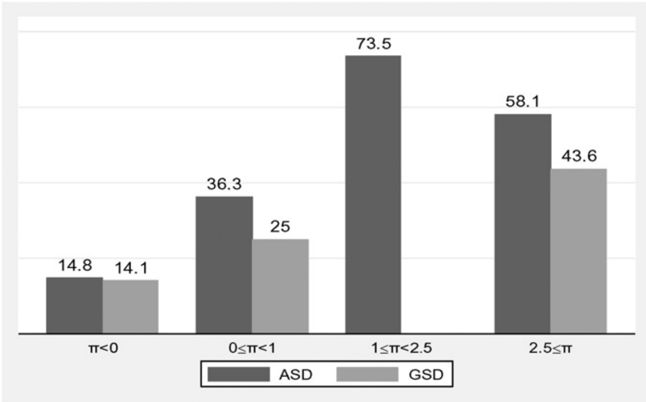

Fig. 1. Aggregate (ASD) and genuine (GSD) inflation state dependence Notes: (i) ASD: The probability of remaining in the same inflation regime two consecutive quarters, (ii) GSD: The probability of remaining in the same inflation regime two consecutive minus the probability of entering that regime from the regime of stable inflation.

As part of inflation persistence may be due to observed and unobserved heterogeneity Fig. 1 depicts the magnitude of GSD as defined in Eq. (6). Results suggest that the probability of inflation persistence is influenced by sizeable GSD effects. For example, 96.8% of the observed persistence in the deflation probability is accounted for by GSD. This means that deflation persistence is due to deflation 'intrinsic' persistence. In other words, it is today's deflation per se that increases the risk of tomorrow's deflation. Moving towards higher inflation regimes, the absolute magnitude of GSD increases. At the same time, however, the ratio of GSD to ASD decreases indicating an increasing influence of other unobserved factors on inflation persistence.

These findings have an important implication regarding the nature of policy responses to price developments and highlight the importance of preventing too low or too high inflation in the first place. In particular, the self-enforcing mechanism characterizing deflation suggests that monetary authorities must be alert to tackle deflationary pressures whenever these arise and certainly before they gather strength. As for policies to deal with inflationary pressures, these should not neglect to account for country heterogeneity and particularities. In other words, the battle to tame inflation may require to accompany ECB's policy with economic policies at the national level (e.g. structural policies). Moreover, the considerable size of GSD suggests that short-term economic policies to stabilize inflation will have long term effects.

This result combined with the better fit of the ordinal models as well as with the fact that the coefficients of the lagged dependent variables in the ordered models are statistically different to each other suggests that inflation persistence is non-linear. Thus, our modelling approach that estimates inflation persistence taking into account the different level of inflation in the origin and in the destination seems reasonable and suitable for the Eurozone.

4. Conclusions

Our starting point was the argument that, inflation persistence in Eurozone varies depending on the level of previous inflation, and thus we need to allow for a non-linear relation between current and previous inflation. To gain greater understanding of this non-linear relationship, we have constructed four inflation regimes and modelled transitions among these using a random effects ordered model, allowing for unobserved heterogeneity and endogenous initial conditions. Our model was estimated using quarterly data from Eurostat for 11 EMU countries over the period 1997:Q1 to 2015:Q3. A series of tests was performed to check the validity of our argument (non-linear inflation persistence versus linear). All tests confirmed that inflation persistence in non-linear.

Results suggest that deflationary episodes are possible but not very likely. If they occur, however, they are less persistent. Inflation persistence increases as inflation rates scale up and move towards ECB's target rates. Interestingly, once inflation enters the high-inflation regime, it shows a strong tendency to remain there as well. So, despite the concerns about the appearance and the possible persistence of deflationary pressures in the euro area, statistically, it is high inflation rather than deflation that records higher persistence and thus posing a risk on the economic prospects of the European economy. Moreover, short-term policies that lead to price stability seem to have a longer effect, since entering the price stability regime the probability of remaining there is 73.5%.

Of course, this is not to say that the risks of deflation episodes are unreal. Neither do we view deflation, which is associated with recession, as an innocuous monetary phenomenon. Negative inflation can become self-enforcing in the sense that, ceteris paribus, falling prices today can themselves increase the risk of falling prices tomorrow. From a monetary policy perspective, these considerations imply that the ECB should remain alert to adjusting its policy tools in response to adverse economic developments.

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The authors are grateful to Costas Kanellopoulos, Kyriakos Neanidis and Apostolis Philippopoulos for their constructive suggestions and comments. The paper has benefited considerably from the advice of two anonymous referees and the editor. The usual disclaimer applies.

Appendix A

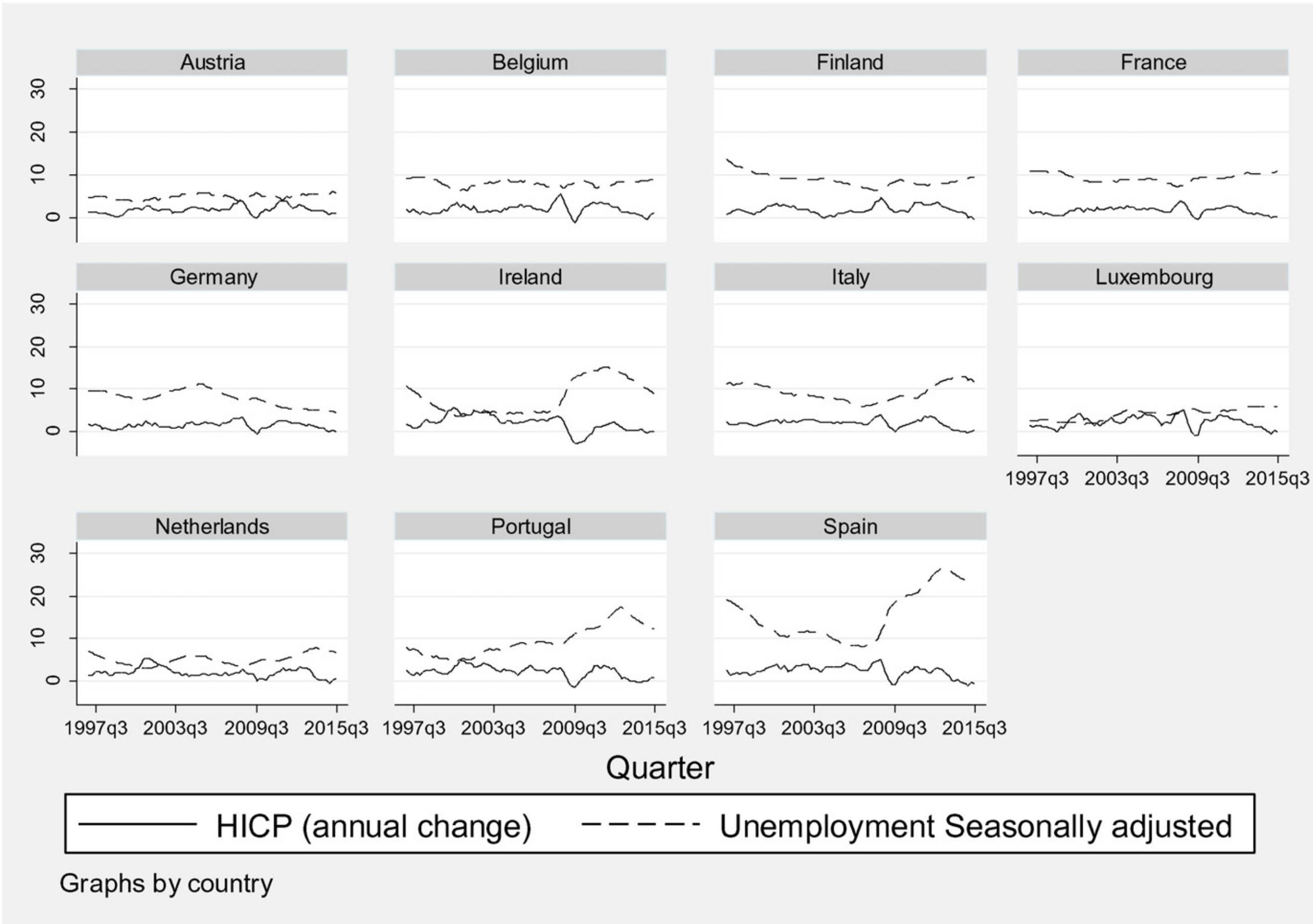

Fig. A1. Evolution of unemployment rate and inflation rate by country, 1997:Q1-2014:Q4

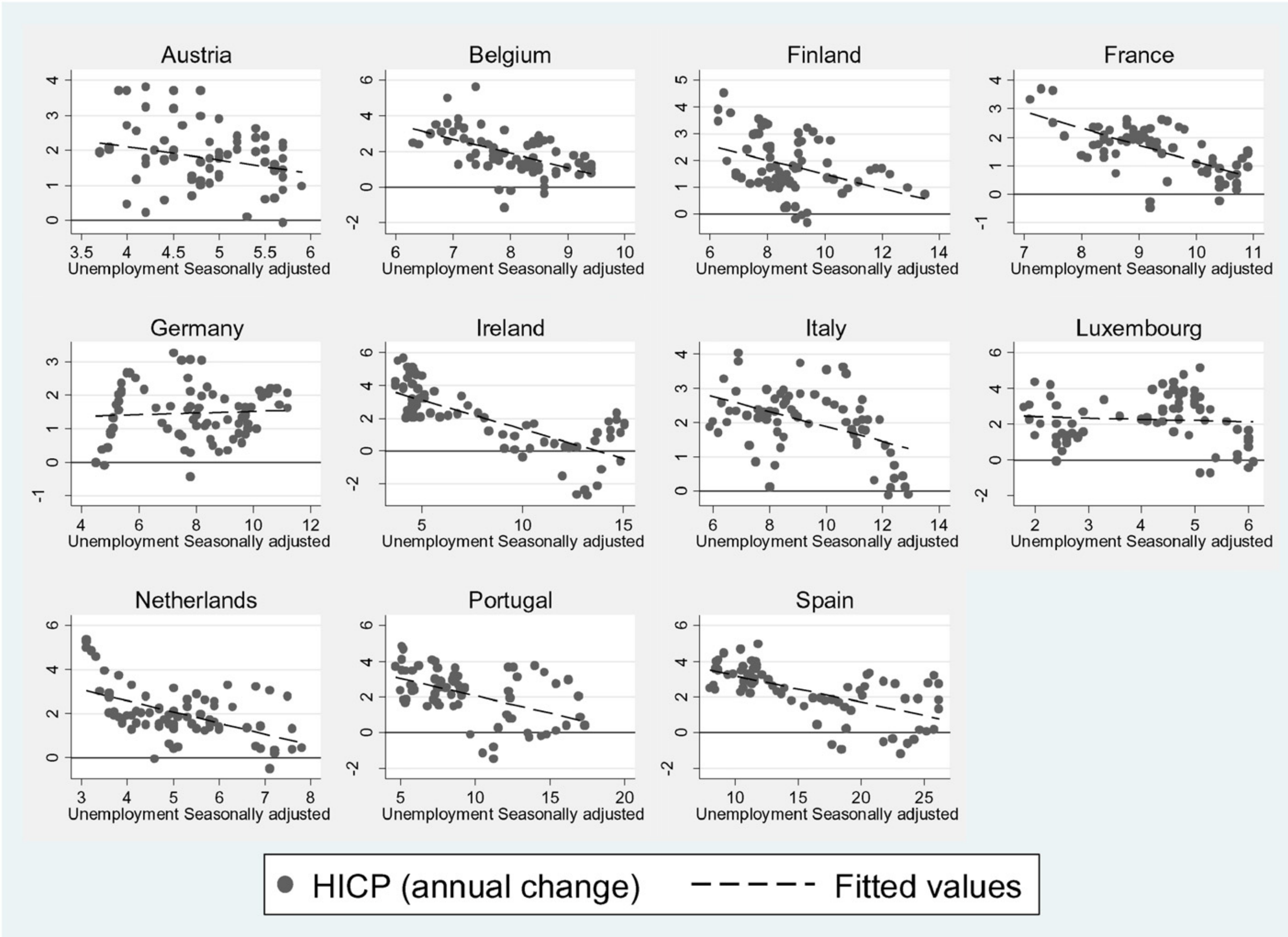

Fig. A2. Correlation between unemployment and inflation rate by country, 1997:Q1-2014:Q4

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