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Merging the Purchasing Power Parity and the Phillips Curve Literatures: Regional Evidence from Italy

by

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Abstract

The main purpose of this paper is to merge together two strands of the literature regarding, either directly or indirectly, inflation: the PPP and the Phillips curve ones. In order to accomplish this task, this contribution applies the tools of the Empirical Growth Literature and of Dynamic Panel Data estimation on a sample of 81 Italian provinces from the year 1986 to the year 1998, exploiting cross-sectional variation to avoid to use instruments not directly connected with the inflation generating process. This research strategy allows to conclude that inflation is characterized by a low degree of persistence and by conditional $\beta-$convergence across provinces. Its most suitable driving variable is the unemployment rate and there are long-term non neutralities at the regional level.

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1 Motivation for the Study

Inflation and prices have been studied so far by different approaches. On the one hand, there is a strand of literature dealing with inflation at a very aggregate level for a single country trying to discover its underlying statistical process as the result of either demand or supply pressures. This literature is usually labelled as the Phillips curve one, either in its traditional form or in the New Keynesian one. On the other hand, there are papers that deal with the issues of the law of one price and of purchasing power parity (PPP), therefore typically considering a number of different countries in the attempt to understand if either the level or the rate of change of prices tends to converge or not.

More recently a number of papers, reviewed below, shifted the focus from between countries differences to within country differences by considering regional inflation differentials. This step can be considerably important for new monetary unions because it can help to understand if their members, passing from the status of countries to that of regions by losing their independence in monetary policy, are destined to long lasting inflation differences or to fast convergence.

The purpose of this paper is to build a bridge between these two strands of the literature in order to help to overcome the limits that they have reached. As it will appear from the literature review that follows, the New Keynesian Phillips curve literature have had problems in finding the right degree of persistence of the inflation rate and what is the independent variable that best helps to explain inflation movements. Furthermore, Abadir and Talmain (2002) showed that output persistence could just be the result of aggregation, so moving from a very aggregate level to the meso level can offer a way to empirically test this claim.

This very change is going to be helpful also in solving the problem of the selection of the exogenous variable driving the inflation rate, otherwise known as the "forcing variable" in the New Keynesian literature. In fact as argued by Becker, Grossman and Murphy (1994), but also by Arellano and Honoré (2001), estimating a model statistically very similar to those estimated below, cross-sectional variability offers invaluable means to identify the model parameters, that time-series studies cannot exploit. Furthermore, moving from the national to the regional level offers a way to reassess a very old controversy, that of a long run vertical Phillips curve. In this context, Hughes-Hallet (2000) argued that by redistributing demand between regions it is possible to minimize the long-run national unemployment rate, even in presence of vertical long-run Phillips curves at the regional level.

Finally, the New Keynesian literature has tried to understand what is the right specification of a micro-founded model for inflation. However, it often neglected the issue if "new knowledge" is any better than "old wisdom". In other words, can the traditional Phillips curve perform any better than the New Keynesian one? This is one of the major questions that I am going to tackle in this paper.

Regarding the PPP literature, it has recently moved some steps from the analysis of between countries to within countries price or inflation differentials in the attempt to test if inflation differentials are really long lasting also in absence of exogenous factors hampering long run price or inflation convergence. The literature have traditionally focused on absolute convergence between inflation rates (prices) in different places, either countries or regions. This is because, if the PPP hypothesis holds, the real exchange rate $\frac{P_f e}{P}$ - where P_f is the foreign level of prices, e is the exchange rate and P is the domestic level of prices - will be a constant, implying, in continuous time, that

$$\frac{P_f'(t)}{P_f(t)} + \frac{e'(t)}{e(t)} - \frac{P'(t)}{P(t)} = 0 \tag{1}$$

where $P_f'(t)$, $P'(t)$ and $e'(t)$ are time derivatives. Under fixed exchange rates $e'(t) = 0$ and $\frac{P_f'(t)}{P_f(t)} = \frac{P'(t)}{P(t)}$, in principle entailing that inflation rates should converge to the same value in different regions (absolute β-convergence). Indeed, in this contribution the PPP hypothesis will not be tested checking for cointegration among regional prices, rather particular care will be devoted to assessing the convergence property of local inflation rates, by estimating a $\beta-$convergence equation (see eq. 6).

To this purpose and as discussed in detailed below, the New Keynesian and Traditional Phillips curves may well offer models of conditional beta convergence (Sala-i-Martin, 1996) alternative to the beta convergence equation that is usually borrowed from the theoretical growth literature, but that has no economic content when applied to inflation.

Moreover, in this context the problems of the selection of the right "forcing variable" and of the assessment of which model fits the data better between the New Keynesian Phillips curve and the traditional one assumes new significance. The literature on inflation differentials highlighted different mechanisms that may explain their long half-life, such as productivity growth differentials between sectors (Balassa-Samuelson effect), inflationary bottlenecks caused by market rigidities unevenly spread across the economy or demand pressures affecting some sectors or regions more than others.

The selection between a "cost based" or an "output gap" Phillips curve should help to understand if the first two stories are more reliable than the latter. If the first model, based on real unit labour costs, fits the data better, then inflation differentials will probably have supply side causes, whereas if the latter does the job better then demand pressures and good market imperfections are likely to be the most important factors. On the other hand, if it is possible to find that the traditional Phillips curve fits the data better, it will have two implications. First regions with a higher unemployment rate lack effective demand, having lower inflation rates than regions with a lower unemployment rate. Second, lower inflation rates will mean higher real interest rates, constraining further aggregate demand.

Furthermore, shifting from absolute to conditional convergence may be particularly interesting, because one of the major finding in the PPP literature, in contrast with the New Keynesian literature, is that inflation does not converge quickly and that inflation half-life is rather long. This high degree of persistence may, in fact, hide conditional convergence and this is one of the hypothesis that this contribution is going to test.

One other major issue for the PPP literature is its inability to explain satisfactorily the fact that inflation usually displays beta convergence but not smooth sigma convergence and the high volatility of the inflation rates, namely their high probability to move from high levels to low levels or viceversa in different time periods. Once noticed these puzzles, it is customary to argue that beta convergence and sigma convergence are linked but do not coincide as sigma convergence depends also on possible disturbances deriving from exogenous shocks (Sala-i-Martin, 1996, Cannon and Duck, 2000 and Bliss, 2000)[1]. Moreover, these very shocks can help explaining the high volatility of inflation rates.

By using a forward looking Phillips curve specification of the model, it is not only possible to shed some more light into the black box of the variance of the exogenous disturbance but also to understand that inflation expectations may play an important role not only in determining the inflation aggregate level but also its cross-sectional variation, helping to explain the wanders of the dispersion of regional inflation rates as well as their volatility. The traditional regional PPP literature could not grasp this point because it focused only on backward looking models.

The rest of this paper is structured as follows. Sections 2 and 3 give a picture of the Phillips curve and the PPP literatures. This review will

[1] In absence of shocks beta and sigma convergence would coincide. Suppose in fact that inflation is generated by AR(1) process $\pi_{it} = \alpha + \beta\pi_{it-1}$ without shocks, π_{it-1} is the rate of inflation in province i at time $t-1$. Suppose also that $E(\pi_{t-1}) = \mu_{t-1}$ and that $Var(\pi_{t-1}) = \sigma^2_{t-1}$. This will imply that $E(\pi_t) = \alpha + \beta\mu_{t-1}$ and $Var(\pi_t) = E[\pi_t - E(\pi_t)]^2 = E\left[\alpha + \beta\pi_{t-1} - \alpha - \beta\mu_{t-1}\right]^2 = E\left[\beta\pi_{t-1} - \beta\mu_{t-1}\right]^2 = \beta^2 Var(\pi_{t-1})$. Hence $Var(\pi_t) < Var(\pi_{t-1})$ if $\beta < 1$. It is the presence of shocks, as correctly stated by Bliss (2000), that makes $\beta-$convergence neither a sufficient nor a necessary condition for σ-convergence.

also offer a way to introduce the models that will be estimated and to further assess the importance of merging the two models under concern. Section 4 describes the dataset and shows some of its features in terms of the distribution and cross-sectional volatility of inflation rates. The aim of this exercise will be to see if the dataset fits the main stylized facts appeared in the literature so far and if the proposed explaining variables display similar patterns to inflation. Section 5 offers a brief methodological review to highlight the reasons underlying the choice of the method of estimation, system GMM, and to offer an introduction to spatial filtering. Section 6 is devoted to the main estimation results and, following Weber and Beck (2003), it will assess the relationship between the average level of inflation and its regional dispersion. Section 7 concludes.

2 The New Keynesian Phillips Curve

Lucas (1976) argued that traditional policy evaluation exercises were flawed by the interaction of policy interventions and economic agents' expectations and spurred a new effort to build solid microfoundations for macroeconomic models. One of the results, regarding monetary issues, was a reinterpretation of the Phillips Curve in the context of the analysis of the rules for monetary policy (see for instance Clarida, Gali and Gertler, 1999). However, the performance of the structural Phillips curve, otherwise known as New-Keynesian Phillips Curve (NKPC hereafter), has raised a not less keen debate than its old counterpart mainly on the ground that the reliability of Taylor-type interest rate rules hinges on good estimates of the parameters of the Phillips curve (Rotemberg and Woodford, 1997, Levin, Wieland, and Williams, 1999 and Roberts, 1998).

Before moving to consider the results of empirical estimates, it is worth briefly considering how the New Keynesian Phillips curve is obtained by standard microfoundations. Building on the New-Keynesian literature about monopolistic competition and differentiated goods (Rotemberg, 1982, Mankiw, 1985, Svensson, 1986 and Blanchard and Kyotaki, 1987), the New Keynesian Phillips Curve can be obtained from the price-setting problem for a profit maximising firm in the context of monopolistic competition. Each firm has a small amount of monopoly power and sets its price by solving the following problem:

$$\max_{P_t(z)} \sum_{i=0}^{\infty} (\theta\beta)^i E_t \left[\Lambda_{t,i} \frac{P_t(z) - MC_{t+i}^n}{P_{t+i}} Y_{t,t+i}(z) \right] \tag{2}$$

$$s.t. Y_{t,t+i}(z) = \left[\frac{P_t(z)}{P_{t+i}} \right]^{-\rho} Y_{t,t+i}$$

where $\beta^i \Lambda_{t,i}$ is the rate at which the firm discounts earnings at time t+i, $P_t(z)$ is the price set by the z-th firm, P_t is the aggregate price index, MC_t^n is the nominal marginal cost, $Y_t(z)$ is the output of the z-th firm, Y_t is the aggregate level of output and θ is the probability a firm has not to reset its price in a Calvo price setting model.

The solution to the problem above gives the optimal price $P_t^*(z)$, which together with the price index, $P_t = \left[\theta P_{t-1}^{1-\rho} + (1-\theta) P_t^{*1-\rho} \right]^{\frac{1}{1-\rho}}$, and after log-linearising gives the New-Keynesian Phillips curve:

$$\hat{\pi}_t = \lambda \hat{mc}_t + \beta E_t \hat{\pi}_{t+1} \tag{3}$$

Under certain conditions one can write

$$\hat{\pi}_t = \lambda k \hat{y}_t + \beta E_t \hat{\pi}_{t+1} \tag{4}$$

where k is the elasticity of the marginal cost with respect to the output gap, λ is a function of the structural parameters, $\hat{\pi}$ is the deviation from steady state of inflation, $\hat{mc}_t$ is the deviation from steady state of the marginal cost and $\hat{y}_t$ is the deviation from steady state of the output gap. The last two equations already show two possible ways to estimate the New Keynesian Phillips Curve, one using as forcing variable a proxy for the real marginal cost and the other by using the output gap[2].

Though the traditional Phillips Curve is backward-looking, the New-Keynesian Phillips Curve (hereafter NKPC) derived from standard microfoundations is purely forward-looking. Rotemberg and Woodford (1997, 1999) found empirical support for this model once allowing a serially correlated error term, however Roberts (1997, 1998 and 2005), Fuhrer and Moore (1995) and Estrella and Fuhrer (2002) found opposite results. The controversial empirical performance of the model induced the literature to take two steps. In the first place, inflation lags were introduced under the hypothesis that some agents form their inflation expectations by looking at its past values. Secondly, the forcing variable was changed from the output gap to the real marginal cost, which, being an unobservable, has been proxied by the real unit labour cost:

[2] For a complete derivation of the New-Keynesian Phillips curve see Gertler (2002).

$$\hat{\pi}_t = \lambda \hat{mc}_t + \gamma_f E_t \hat{\pi}_{t+1} + \gamma_b \hat{\pi}_{t-1} \tag{5}$$

where $\hat{mc}_t$ is deviation of the real marginal cost from steady state. The model above was labelled the "hybrid model" and successfully estimated by both GMM and a two-step distance-VAR method (Gali and Gertler, 1999, Gali, Gertler and López-Salido, 2001, and Sbordone, 2001).

This model specification attracted a considerable attention. Guerrieri (2001) showed that the estimate of Gali and Gertler (1999) are not robust to alternative normalizations of the moment conditions, though their normalization appears to be superior to other alternatives in a Monte-Carlo experiment, and that using Taylor-style contracts instead of Calvo ones the share of backward looking firms increases dramatically. Roberts (2005) found that the result of tiny relevance of backward looking pricing behavior of both Gali and Gertler (1999) and Sbordone (2002) hinges on proxying the marginal cost by real unit labour cost and that exactly this fact allows an interpretation of their result in the light of a traditional aggregate Phillips curve. Jondeau and Le Bihan (2005) - by using both the GMM and the ML estimators on data from Europe, the US, the UK, France, Germany and Italy - found that the NKPC has to be augmented with additional inflation lags and leads, that there is a large fraction of backward looking firms and that the most suitable forcing variable for the UK and the US is the marginal cost while for the other countries it is the output gap. Dupuis (2004), by contrast, finds that for US data an NKPC with the output gap provides better inflation forecasts than both the NKPC with the marginal cost and that with polynomial adjustment costs as proposed by Kozicki and Tinsley (2002 a, b). On the other hand, Gagnon and Kahn (2005) by considering US, Canadian and Euro area data and tackling aggregation issues using both a Cobb-Douglas production function and a CES technology obtained the following results: ignoring aggregation issues gives implausible price stickiness and the marginal cost is a better forcing variable than the output gap when using a CES technology instead of a Cobb-Douglas one, particularly for Canada and the Euro area.

Kurmann (2005) argued that the results obtained by Gali and Gertler (1999) and Sbordone (2002) hinge on both the calibration of the structural pricing equation implied by the Calvo model and the reliability of a reduced form forecasting process for the real marginal cost and that both the assumptions are questionable. Sbordone (2005), giving a more general interpretation of her previous work, stressed that the issue of uncertainty can be tackled with her methodology while the preliminary stationarity-inducing transformations, the size of the model and the lag length are still open questions. Cogley and Sbordone (2005) addressed the time invariance of the deep parameters by using a two-step method: first they estimated a VAR with drifting parameters, then taking as given the parameter estimates of the unrestricted VAR they estimated the NKPC by minimising a quadratic function of the restrictions that the theoretical model imposes on the reduced form. They found that it is possible to reconcile a time-drifting VAR with a constant parameter NKPC confirming its structural stability.

Neiss and Nelson (2002) argued that the better performance of "cost based" with respect to "output gap based" Phillips curves is mainly due to the inappropriateness of the most widespread output gap measures for DSGE models, not considering that in this context potential output is affected by real shocks and that output gap is not a business cycle indicator rather than one of nominal rigidities. Consequently, they propose a new output gap measure, as an approximated infinite lag sum of preference and technological shocks, consistent with DSGE models. They find that a NKPC based on this output gap proxy fits Australian, UK and US data well.

Rudd and Whelan (2005) showed that time-series GMM is not able to distinguish between a forward-looking and a backward looking Phillips Curve due to specification errors and to the apparent irrelevance of both forward-looking expectations for the current rate of inflation and of past inflation rates to present values of the forcing variable. Lindé (2005) also pointed out that estimating the model by non-linear least squares, in order to avoid the problems emphasized by Rudd and Whelan (2005), provides evidence against the New-Keynesian Phillips Curve, that time-series GMM estimates are biased and that estimating the model with full information maximum likelihood seems a more attractive procedure.

Gali, Gertler and López-Salido (2005) replied that the closed form used by Rudd and Whelan (2005) does not allow a satisfactory evaluation of the hybrid model because based on the null hypothesis of a purely forward-looking inflation generating process, that the results obtained by Lindé (2005) with NLS are plagued by endogeneity of the independent variables and that the FIML estimator assumes that the econometrician has a good deal of knowledge regarding the true model of the economy. However, the NKPC did not manage to survive a sensitivity analysis of the instrument set and a battery of specification tests in Bårdsen et al. (2005).

Both Roberts (1995 and 2005) and Sbordone (2005) pointed out that the cost based version of the NKPC leaves unexplained the movements of the marginal costs and the former that, given labour hoarding, very often marginal cost and unit labour costs are not thought to move together. In order to overcome this shortcoming, Blanchard and Gali (2005), on the ground of Trigari (2004), proposed to use as forcing variable the unemployment rate, offering microfoundations for the traditional Phillips curve augmented with expectations.

3 The PPP Literature

3.1 From Time Series to Cross Section Studies

The PPP literature was subject in the past to a series of methodological shifts. Early studies were mainly concerned with time series data, attempting to compare the percentage changes in bilateral exchange rates with inflation differentials. Their failure to find evidence in favour of the PPP hypothesis spurred an attempt to collect data for very long time series, in order to dispel the doubts concerning the low power of tests used to detect convergence towards a long-run equilibrium. The results this time were in favour of the PPP hypothesis, with deviations from PPP having an half-life of around 4 years (Frankel and Rose, 1996).

This solution to the problem was not without its own shortcomings because long time series are subject to potentially serious structural shifts. Therefore, Frankel and Rose (1996) proposed to consider a cross-sectional approach in order to achieve the necessary degree of variation in the data to obtain enough powerful tests. Building on this methodological change, Imbs et al. (2005) showed that, considering sectoral heterogeneity, the half-life of the real exchange rate may fall from the "consensus view" of 3-5 years to eleven months.

A similar concern could be raised regarding the NKPC literature, because the debate between Lindé (2005) on one side and Gali, Gertler and López-Salido (2005) on the other showed two important facts. First, the choice to use an instrumental variable estimator or not can have dramatic consequences on the very sign of the parameters under concern. Second, to obtain results in favour of the New Keynesian Phillips curve it is necessary to use instruments that are not directly connected with the specified stochastic process for inflation, allegedly on the ground that they are used by agents to form their expectation, though the way these expectations are formed is not specified or investigated at all[3].

Furthermore, using a panel data approach may help not only to overcome the problem of heavy instrumentation, but also that of potential

[3] For instance, the GMM estimation performed in Galì and Gertler (1999) uses as instruments four lags of inflation, the labor share of income, the long-short interest rate spread, the output gap, wage inflation, and commodity price inflation. In Galì, Gertler and Lopez-Salido (2001) a more parsimonious instrument set was chosen, including not only four lags in the inflation rate but also two lags in the real marginal cost, in de-trended output and wage inflation for both the cost based and the output gap based NKPC.

structural shifts in datasets that usually start in the sixties and therefore include different monetary policy regimes and, potentially, different way for agents to form their expectations.

One other major finding of the PPP literature, is that the catching up process of countries with a low level of prices towards those with a higher one cannot offer a completely satisfactory explanation for inflation differentials between countries and that factors other than price convergence can explain most of the cross-country inflation differences (Rogers, 2001). This points to the potential benefits that can be reaped by merging the PPP literature with the Phillips curve one, because this can help to better understand what these factors are.

3.2 The Regional PPP Literature

The presence of factors hampering the adjustment of relative prices (inflation rates) spurred researchers to move to consider regional datasets, as another way to predict if countries joining a monetary union were doomed to go through a long lasting adjustment process or destined to a quick smooth convergence. Among the main factors hampering relative price adjustment it is possible to list: a) tariff barriers; b) non-tariff barriers; c) nominal exchange rates failing to adjust to relative price-level shocks; d) market imperfections allowing firms to apply different price policies in different countries; e) costs in adjusting prices; f) transportation costs hampering arbitrage between different countries; g) the presence of non traded goods, for which arbitrage is impossible (Cecchetti et al., 2002).

Other explanations that have been offered by the literature to explain price (inflation) differentials are: i) a positive correlation between the level of income and the level of prices, implying that catching up regions or economies should experience positive inflation differentials; ii) macroeconomic disequilibria, whereby it is not said that all the regions within a country experience the same demand pressures; iii) even in presence of the same demand pressures there might be different market rigidities, implying stronger or weaker inflationary bottlenecks (Alberola, 2000). It is worth noting that an output gap based NKPC should be able to verify if these hypotheses find any support in the data, because inflation should be found to positively depend on the difference between actual output and the natural output, which is larger the stronger are demand pressures[4] or the greater are market imperfections[5].

One of the major studies of "within countries" price differences is

[4] That should be sizeable for catching up economies.

[5] According to microfounded models, without market imperfections the difference between actual and nominal output would not exist.

Cecchetti et al. (2002) who analysed a dataset of the price indexes of 19 major US cities from 1918 to 1995 with the following characterization of the data:

$$\Delta q_{it} = \mu_i + \theta_t + \beta_t q_{it-1} + \sum_{j=1}^{k_i} \gamma_{ij} \Delta q_{it-1} + \epsilon_{it}$$

where q_{it} is the log of the price level of city i at time t, μ_i is a city specific constant, θ_t is a time specific constant and γ_{ij} are the lag coefficients in the process characterizing q_{it}. They found that, considering the whole sample, relative price adjustment has an half-life of 8.5 years[6]. They proposed three explanations for such a slow convergence: distance - on the account that the price differential between two cities is larger the farther the two cities are -, different adjustment costs for small and large deviations and non traded goods. Remarkably, they did not manage to find any statistical support for these three explanations. They also could not test if the real wage or productivity differentials could affect their results due to data constraints. Here it is possible to perform a similar exercise but with data on a greater number of variables and it is possible to see that in fact high persistence may hide conditional convergence.

Parsley and Wei (1996) analysed a quarterly data set including 51 final tradable and non-tradable goods and services from 48 cities from 1975 to 1992. They find that distance, proxying for arbitrage costs, does affect the size of price differences and its convergence rate, therefore the more two cities are distant the more price differentials are variable and wide and the longer they take to converge. A similar role for distance was found by Engel and Rogers (1996a). Besides the role of distance, Parsley and Wei (1996) highlighted that prices of tradable goods converge faster than non-tradable ones, in contrast with the results found by Cecchetti et al. (2002).

Finally, Weber and Beck (2003) analysed a panel of 77 European regions from 1991 to 2002 using monthly data and a very similar model to Cecchetti et al. (2002) but for inflation instead of the price level. They find that: i) regional inflation rates do not display smooth sigma convergence; ii) they do display a lot of internal volatility - whereby regions with a high inflation ranking in the present may have a low one in the future; iii) there is a positive relationship between regional inflation dispersion and mean which can allow central banks to decrease the average inflation down to 1% without having a sizeable percentage of regions to enter deflation; iv) mean-reversion takes place at a slow

[6] Though it displays a sizeable subsample variability, reaching the value of nearly 53 years in the period from 1956 to 1975.

pace, that is inflation half-life can be rather long, ranging from 0.5 to 75.1 years for different sub-samples.

3.3 The Balassa-Samuelson Hypothesis

The effect of productivity growth differentials is one of the major topic of the literature of inflation differentials within countries and it can be traced back to the pioneering studies of Balassa (1964) and Samuelson (1964). The Balassa-Samuelson model hinges on the distinction between a tradable and a non-tradable sector. In the former, prices are determined on the international market and a strong productivity growth takes place, whereas in the latter prices are determined locally and productivity growth is weak. The two sectors are assumed to have the same wage which is linked to productivity growth in the tradable sector. Therefore, the faster productivity grows the stronger is the wage push and the greater is inflation in the non-tradable sector, where productivity growth cannot absorb the increase in the wage bill (Alberola, 2000). In other terms, there should be a positive relationship between productivity growth and inflation.

However, the assumptions of the same wage across different sectors, or even of a constant ratio between wages of different sectors, seems very restrictive even supposing the presence of large unions and taking into account that wage bargaining takes often place along sectoral lines[7]. It seems advisable to consider as regressor not productivity growth, but the real unit labour cost, like in the New Keynesian Phillips Curve, because it is known that individual wage bargaining is very widespread and can lead to regional differentials in nominal wages - that in Italy, for instance, are known to be around 20% (Vaona, 2003). Therefore, the real unit labour cost may more easily account for the opposing forces of productivity growth and wage costs on inflation.

This could be one of the reasons why the Balassa-Samuelson hypothesis was rejected in Spain at the provincial level (Alberola and Marqués, 1999). Further evidence against the Balassa-Samuelson effect was found also by considering countries instead of regions. For instance, Rogers (2001) points to the fact that in the year 2000, Portugal, Ireland and to some extent Greece had both high inflation and productivity growth, but this was not so for Austria and Spain. Honohan and Lane (2003) detected a negative cross-sectional correlation between productivity growth and inflation in the period 1997-2001.

[7] For instance in Italy, during the nineties, the wage purchasing power was stable in the manufacturing and trade sectors and declining in agriculture and other service sectors (Birindelli, D'Aloia and Megale, 2003).

3.4 Inflation Differentials Within EMU

Another strand of the literature to be considered is that concerning inflation differentials within EMU. Busetti et al. (2006) could not reject the hypothesis of convergence between 1980 and 1997, but they could between 1998 and 2003. In this period the formation of three clusters was detected: a low inflation group - including Germany, France, Belgium, Austria, Finland and Luxemburg –, a high inflation one - including Spain, the Netherlands, Greece, Portugal and Ireland – with Italy, constituting the third group, staying in between. These clusters resulted to be driven by the dynamics of unit labour cost, but also by those of productivity and mark-ups, stressing the importance of conditional $\beta-$convergence to understand inflation differentials.

It is possible to reach similar conclusions considering Honohan and Lane (2003) who found that inflation differentials in Europe can be explained by the different impact the Euro devaluation had on the member states of the monetary union, together with differences in the output gap. Angeloni and Ehrmann (2004) stressed the role of three mechanisms to explain inflation differentials in a monetary union. First of all, high inflation countries will have a lower ex-post real interest rate and, if inflation expectations are country specific, even a lower ex-ante real interest rate (dis-equilibrating mechanism). Secondly and in contrast, a high inflation country will tend to lose price competitiveness, reducing its aggregate demand and output (re-equilibrating mechanism). Thirdly, inflation stickiness will help to propagate inflation differentials, due to the inability of the different countries to adjust inflation rates quickly.

4 The Dataset and Its Features

The dataset covers inflation rates in CPI, unemployment rates, the value added and the real unit labour cost for 81 Italian provinces at annual frequency from 1986 to 1999. Data about inflation rates are produced by the Italian national statistical office (ISTAT) and they are the basis for the computation of the national index. Also data about unemployment rates are produced by the Italian statistical office, as well as the data about employment that are used to compute the real unit labour cost. The Tagliacarne Institute releases every year data for the value added at the provincial level, whereas the wage bill was proxied by the average wage for each province resulting from administrative data released by the Italian Pension Institute (INPS)[8].

[8] The source of the data for inflation, unemployment, employment and the value added is a valid guarantee against measurement error. Furthermore, these very data are widely used by local policy makers and workers' and firms' associations to assess

Italy counts 103 provinces, but especially for many Southern provinces it is impossible to recover the data, so 81 provinces are actually covered. Out of 868 observations, 448 (46%) come from the North, 204 (24%) come from the Centre and 216 (30%) come from the South. Though these percentages are not so far from one another, this could entail a self-selection problem: estimates may be distorted by the predominance of Northern provinces. However, as showed later, I run separate regressions for the North and the Centre-South and I could not reject the null hypothesis of equality between them. Furthermore, the signal to noise ratio is stronger in the Centre-South than in the North, allowing to think that the addition of more Southern provinces would confirm the results achieved in this contribution. In the empirical estimates below, when dealing with regional disparities, in order to assure comparability in the sub-sample sizes I pooled together the South with the Centre and I called this sub-sample "South". Having two comparable sub-samples in terms of cross-sectional units is important due to the finite sample properties of the GMM panel estimator (Alonso-Borrego and Arellano, 1999).

In the economic growth literature, convergence has not been studied only by making use of regression analysis but also resorting to a distributional approach after Quah (1996) and Quah (1997) among others. In this way, it was possible to highlight that economic convergence is mainly taking place in clubs and that the distribution of the growth rates is characterized by "emerging twin peaks", that is by a polarization between a high income and a low income club[9]. By applying these tools to inflation or price convergence it is possible to gain similar insights, as showed by Gluschenko (2004) and Weber and Beck (2003), and to provide valuable information to point out new stylized facts for regional inflation differentials, useful when selecting the best specification for a parametric model. In this contribution, I mainly rely on Gaussian Kernel Estimators using Silverman (1986) optimal smoothing bandwidth, given that different kernel functions have been showed not to affect estimation results in a significant way (Pagan and Ullah, 1999). For sake of brevity, I show here only the stochastic kernel for inflation. Its contour plot together with the stochastic kernels and the contour plots of the other variables are available from the author upon request..

the state of local economies. Regarding the measure of the real unit labour cost, I aggregated the data by NUTS 2 regions - the present study is carried out for NUTS 3 regions - and I computed the mean and the standard deviation of the resulting variable obtaining respectively 0.50 and 0.06. Once taking from the regional accounts data about the ratio of regional wage and salaries over total households' income I obtained a mean of 0.55 and standard deviation of 0.05.

[9] For a recent survey, see Durlauf and Quah (1999).

Figure 1 shows the standard deviation and the average of the provincial inflation rates for each year. It appears clear that the years from 1987 to 1998 have been years of deflation, though not of smooth deflation. In fact, the average inflation rate moved up and down until 1991 to decrease markedly between 1991 and 1994, experiencing a new peak in 1995 and falling markedly again until 1998[10].

It is also worth stressing that Figure 1 highlights two features of the inflation generating process that has been found also in Weber and Beck (2003) and in the literature regarding inflation differentials in EMU. First, the lower is the average inflation rate the smaller is its geographical dispersion, and, second, though inflation dispersion decreases, it does not do it smoothly alternating periods of sigma convergence to those of sigma divergence.

Further insights can be gained by considering the stochastic kernels of the variables under scrutiny. Figure 2 shows the stochastic kernel for inflation. The fact that the distribution is parallel to the $t - \tau$ axis means that inflation passed from high and well dispersed values to low and more concentrated values. Inspecting the contour plot it is possible to divide the provinces into two groups, one that had medium or low inflation rates in 1987 and kept them in 1998 and one that had high inflation rates in 1987, but managed to underbid the first group by the end of the nineties.

There is evidence that this dichotomy reflects a geographical pattern[11]. Figure 3 displays the actual data and it has the provincial inflation rates in 1987 on the vertical axis and those in 1998 on the horizontal one. Again it is possible to see two groups: the first mainly in the North having low inflation rates in both the periods and the second group, mainly in the South, having high inflation rates in 1987 and low inflation rates in 1998. There is clear evidence that the Southern provinces underwent a period of intense deflation that led them to have lower inflation rates than their Northern counterparts.

Let us move to consider the candidate forcing variables. The estimated stochastic kernel for the log of the value added appeared to be clearly placed along the diagonal of the Cartesian plane, indicating a high degree of persistence in the distribution. The presence of three peaks in the stochastic kernel showed that Italian provinces can be divided into three groups: those that had a low value added both in 1987 and in 1998, those that had a medium value added in both the years and

[10]Note that 1987 and 1998 are respectively the first and the last year of estimation due to the presence of one inflation lag and of inflation expectations in the model.

[11]The groups showed in Figures 3-6 were obtained by inspecting the contour plots of the stochastic kernels of the variables under analysis.

those that had a high value added in both the years. Inspecting Figure 4 it is possible to conclude that the first group is mainly composed by Southern Provinces, the second by Northern Provinces and the third by four outliers, two in the North and two in the South.

A similar geographic divide emerges by considering the data regarding the unemployment rates. Again the position of the estimated stochastic kernel showed a highly persistent distribution with three groups of provinces: the high unemployment ones, the medium unemployment ones and the low unemployment ones. Figure 5 shows that the last group is mainly composed by Northern Provinces, the second by a mixture of a minority of Northern provinces and a majority Southern ones and the first only by Southern provinces.

When considering the log of the real unit labour cost, the contour plot of the stochastic kernel did not show a stable distribution as in the previous two cases: the provinces can be divided into three groups: two groups remained stable around respectively a medium real unit labour cost and small one, while part of the third group shifted from a high real unit labour cost to a medium one. Figure 6 shows that the low real unit labour cost group is composed by an outlier in the North, the medium one is mainly made by Northern provinces and the partially-shifting group by Southern provinces.

After a detailed look at the data, it is possible to make some inferences. First of all, the inflation generating process displays a lot of variability across time in terms of shifts in both its average and dispersion (Figg. 1 and 2). It is not possible to trace these shifts back to changes in the candidate explaining variables because they do not display the same degree of volatility. Therefore, it is clear that a model attempting to estimate the inflation generating process should include also inflation lags as done by both the PPP and the Phillips Curve literatures, but also inflation leads as done only by the Phillips Curve literature; a so high degree of variability cannot be reduced only to persistence or exogenous shocks. By contrast inflation expectations are very likely to play a role, as a volatile factor causing inflation to jump up and down.

However, from the analysis above, it also emerged that there has been at least another factor that has shifted inflation rates downward more in the South than in the North and this factor is unlikely to have been only inflation expectations. By contrast there should be a variable explaining the geographical concentration of the deflationary process. It was showed above that among the candidate "forcing variables", the log of the value added and especially the unemployment rate signalled that the South was lagging behind the North for all the period under consideration, while also experiencing, in part, a greater reduction in real unit labour cost when compared to the North. Therefore, all the three forcing variables can potentially offer explanatory power to a model and regression analysis is strongly called for to select the one that best fits the data.

5 Estimation Methods

The model to be estimated is the following:

$$\hat{\pi}_{it} = \lambda \hat{x}_{it} + \gamma_f E_t \hat{\pi}_{it+1} + \gamma_b \hat{\pi}_{it-1} + \mu_i + \theta_t + \epsilon_{it} \tag{6}$$

where, as stated above, $\hat{\pi}_{it}$ is the inflation rate and $\hat{x}_{it}$ is the forcing variable, either the log of value added, the unemployment rate or the log of the real unit labour cost. In the case $\hat{x}_{it}$ was the the unemployment rate and the hypothesis $\gamma_f + \gamma_b = 1$ holds, then θ_t would be the aggregate non-accelerating-inflation-unemployment-rate. Also for this reason it is important to test if $\gamma_f + \gamma_b = 1$. No exchange rate term appears in (6), because the exchange rate between regions is fixed (see equation 1).

In general, a β convergence equation can be written as follows

$$y_{it} = \alpha + \beta y_{it-1} + \epsilon_{it} \tag{7}$$

where y_{it} is the variable of interest, α is the common factor and ϵ_{it} is a stochastic error. In the empirical growth literature it has become customary to estimate a different equation derived from microfounded economic models:

$$\log\left(\frac{y_{i,t}}{y_{i,t-1}}\right) = a - \left(1 - e^{-\beta}\right) \cdot \log\left(y_{i,t-1}\right) + \epsilon_{it} \tag{8}$$

Equation (8) , however, does not have any economic content when applied to inflation and it seems better to resort to a model specification similar to (7), and therefore as in (6) , which was originally adopted by scholars studying convergence of heights of different generations or of firm sizes (Hart, 1995).

As it is possible to see, (6) requires an instrumental variable estimator not only because of the lag in inflation that is correlated with the error component accounting for spatial heterogeneity (μ_i), but also because of the inflation lead which is endogenous because depending not only on μ_i but also on ϵ_{it}. The insertion of time effects will help to tackle the issues of a time varying natural level of output, overcoming the critique of Neiss and Nelson (2002), and of serial correlation in the residuals, in order to have consistent estimates. It will also allow to capture the effect of aggregate common factors. μ_i, instead, will account for spatial differences in the steady state level of the forcing variable.

It is necessary to make two further points about (6). First, it is a model of conditional beta convergence as no constant common factor appears. In the context of regional inflation differentials this means that regional inflation rates will not converge to a unique aggregate value, but to different values depending on the conditioning variables, unless there exists a vertical long run Phillips curve even at the regional level. Indeed and in the second place, conditional beta convergence is deeply connected with the existence of a long-run Phillips Curve at the regional level, because the long-run inflation rate will depend on the regressor $\hat{x}_{it}$, only if $\gamma_f + \gamma_b \neq 1$.

I use the system GMM estimator proposed by Blundell and Bond (1998), which is known to outperform the GMM estimator proposed by Arellano and Bond (1991) in finite samples. Arellano and Bond (1991) estimator is based on first differencing the model and then using past levels of the involved variables as instruments. On the other hand, the system GMM estimator is based on keeping into consideration both the first differenced equation and the equation in levels and on using as instruments for the former ones the past levels and for the latter ones the past first differences[12]. As a first step, I used as instruments all the available lags of inflation, both in differences and in levels, and the current level of the forcing variable. However, as it will appear later, I also reduced the number of instruments to check for stability of the parameter estimates. I chose not to use other variables in the instrument sets, as in Gali and Gertler (1999), because, relying on Arellano and Honoré (2001) and Becker, Grossman and Murphy (1994), I exploited cross-sectional variability to identify parameters, keeping the instrument set to a minimum size, given also the problems of bias that heavy instrumentation can imply for the GMM estimator (Ziliak, 1997).

Therefore the orthogonality conditions used in this study are as follows

$$E\left(\hat{\pi}_{i,t-s}\Delta\epsilon_{it}\right) = 0 \text{ for } t = 3, ..., T \text{ and } s \geqslant 2 \tag{9}$$

$$E\left(\hat{x}_{it}\Delta\epsilon_{it}\right) = 0 \text{ for all } t \tag{10}$$

$$E\left(\epsilon_{it}\Delta\hat{\pi}_{i,t-1}\right) = 0 \text{ for } t = 4, ..., T. \tag{11}$$

[12]It may be possible to argue that being GMM an extension of 2SLS, one may not identify γ_f and γ_b, because he is anyway supposing a model for expectations given that they are correlated with the instruments. However, this is a very restrictive interpretation of the GMM estimator, whose underlying theory was especially devised to estimate directly the effect of expectations on current variables. Furthermore, to impose the restriction $\gamma_f + \gamma_b = 1$ as an identification device even though the data reject it, as in the present application, entails approaching the data with a very strong prior, even though this may not be justified (Mankiw, 2001).

ϵ_{it} can be rationalized as an expectational error (Gali and Gertler, 1999), so the first and the third conditions only imply that past level of inflation or a function of theirs, such as their first difference, are not correlated to expectational errors. The second equation, instead, implies thinking to the forcing variable as exogenous to the inflation generating process, which is consistent with the Phillips curve tradition.

It is worth recalling that after first differencing the residuals of the model above will assume the following form:

$$\nu_{it} = \epsilon_{it} - \epsilon_{it-1} \tag{12}$$

Therefore, absence of serial correlation in the original model, necessary for the validity of (9), will be detected by finding first order negative serial correlation and no second order serial correlation in the first differenced residuals, ν_{it}[13].

It is also worth stressing that in the dynamic panel data literature, the model error is customarily assumed to be identically and independently distributed (Baltagi, 2003), an assumption that would clash with detecting spatial correlation in the residuals. In order to overcome this problem, after testing for spatial correlation in the data, I used the Griffith's eigenfunction decomposition approach to spatial filtering as recommended by Getis and Griffith (2002). This procedure is based on regressing both the dependent and the independent variables on those eigenvectors of (13) with a Moran's I statistic greater than 0.19 (roughly the margin of error threshold value):

$$\left(\mathbf{I}_{NT} - \frac{\mathbf{1}_{NT}\mathbf{1}'_{NT}}{NT}\right)(\mathbf{I}_T \otimes \mathbf{W})\left(\mathbf{I}_{NT} - \frac{\mathbf{1}_{NT}\mathbf{1}'_{NT}}{N}\right) \tag{13}$$

where N is the cross-sectional dimension of the dataset, T is the time dimension of the dataset, $\mathbf{W}$ is an $N \times N$ binary spatial contiguity matrix, NT is the number of observations in the sample, $\mathbf{I}$ is the identity matrix and $\mathbf{1}$ is a vector of ones. By spatial contiguity matrix, it is meant a matrix whose elements, corresponding each to a pair of observations, are equal to one for observations belonging to contiguous regions and zero otherwise (Anselin, 1988).

The Moran's I statistic of a variable y is given by:

$$I^* = \frac{N\left(y'\mathbf{W}y\right)}{S_0\left(y'y\right)} \sim N(0,1) \tag{14}$$

where S_0 is a normalizing factor,

[13] For a deeper understanding of this issue, the reader is referred to Arellano and Bond (1991).

$$S_0 = \sum_{i=1}^{N}\sum_{j=1}^{N} w_{ij} \tag{15}$$

and w_{ij} is the element of the i-th row and j-th column of $\mathbf{W}$. As showed by Anselin and Kelejian (1997), I^* has an asymptotic normal distribution[14]. The meaning of I^* is that the larger is the ratio between the weighted and the unweighted sums of the squares of the elements of y and the stronger is spatial correlation.

It is also worth stressing that spatial filtering assumes a particular relevance within this context due to the findings of the PPP literature that regional differentials in the level of prices are less and less connected the greater is the distance between the two locations they belong to (Engel and Rogers, 1996b and 2001). This also implies that if the path followed by the level of the prices is less correlated the further is the distance between locations the less correlated will also be the inflation rates. In fact, the first column of Table 1 shows that, inflation, but also the other variables considered in this contribution, display strong spatial correlation, that is observations closer in space are more correlated than those further away.

6 Estimation results

The results of the filtering procedure are showed in Table 1. Though it was particularly successful for inflation and unemployment and less effective for the logs of value added and of real unit labour costs, it determined a dramatic decrease in the absolute value of the Moran's I statistic for each variable. It is therefore possible to proceed to estimation on the ground that the dependent variable is not spatially correlated and that the GMM method assumes the errors of the model to be independently and identically distributed, but it does not make any assumption on the regressors (Baltagi, 2003). Therefore, the crucial issue when evaluating estimates will be if the residuals display spatial correlation or not, an issue that will be assessed by appropriate testing.

Table 2 shows the estimation results. The Sargan test as well as the Arellano and Bond test for first order and second order serial correlation support the model. Over-identifying restrictions and the null of no second order correlation of the differenced residuals are not rejected, while,

[14] An alternative spatial filtering procedure was proposed by Getis and Ord (1992) and used to detect beta regional growth convergence in Europe by Badinger et al. (2004). The two procedures have been reviewed by Getis and Griffith (2002) and showed to be equivalent. I used Griffith's eigenfunction decomposition approach as computationally more appealing.

as expected, there is evidence of negative first order serial correlation in the differenced residuals. The Moran's I statistic for the residuals does not reject the null of no spatial correlation for any of the three models, testifying the success of the filtering procedure carried out above. For all the models presented, the coefficient of the inflation lead is comparable to the aggregate estimates obtained by Jondeau and Le Bihan (2005) for Italy. However the coefficient of the inflation lag is much smaller, supporting the view of Imbs et al. (2005) according to which persistence is a property of aggregate time series that disappears once moving to analyse disaggregated ones.

Columns 1, 2 and 3 of Table 2 give information about which forcing variable best fits the data. However, neither the log of the value added, nor the log of the real unit labour cost are significant at the 5% level though their coefficients have the expected sign and magnitude. By contrast, the coefficient of unemployment has the expected negative sign and it is significantly different from zero at the 5% level. Therefore, both the output gap based and the cost based NKPC break down, whereas the traditional Phillips curve seems a more promising model to depict the inflation generating process at the regional level in Italy.

Regarding the restriction of the vertical long run Phillips curve (namely that the sum of the coefficients of inflation lags and leads to be one), a Wald test strongly rejected it reporting a value of 27.22 with p-value of 0.00.

Furthermore, the significance of the time dummies imply that the long run relationship between inflation and unemployment is not a stable one. Though in presence of a vertical long run Phillips curve time dummies could be interpreted as changes in the non-accelerating inflation unemployment rate, here given the rejection of the hypothesis $\gamma_f + \gamma_b = 1$, they can be interpreted as exogenous shifts in the long-run inflation-unemployment trade-off. Figure 11, based on the model specification of column 4, displays the long-run inflation unemployment trade-off for different years: it appears clear that the largest outward and inward shifts took place respectively in 1989 and in 1998, the years with the higher and the second lower average inflation rates (Fig. 1)[15].

In columns 4 and 5, I respectively improved the model specification by discarding the non significant time dummies and I checked for pa-

[15] It is worth recalling that in order to avoid the dummy variables trap, when estimating the model with a constant, it is necessary to drop one of the time dummies, so one cannot directly distinguish what is the impact of the constant and what is the impact of possible shocks in the year without dummy. In order to accomplish this task, I considered as intercept of the long-run unemployment inflation trade off the average of the value of the coefficient of the dummies and as shifts idiosyncratic to specific years the deviations from this average.

rameter stability by running a two-step system GMM estimation with Windemejir (2005) small sample correction. Results display remarkable stability and the significance of unemployment further increases.

Further robustness checks where carried out by considering sub-sample stability. Columns 6 and 7 show estimation results considering respectively the years before and after 1993. Point estimates are slightly different. In fact performing a joint Wald test for inflation parameters, the null of equality between the coefficients of the two sub-period models was rejected at the 5% level (though not at the 1% level). The test statistic, distributed as a χ^2 with 2 degrees of freedom, returned a value of 7.66 with a p-value of 0.01. For the parameter of unemployment, instead, the null of equality across the two sub-periods could not be rejected (the Wald test returned a value of 0.99 with a p-value of 0.32). On the one hand, for unemployment this is clearly a case where pooling across time reduces the coefficient variance: trading in some bias for a reduction in the noise is a desirable step (Baltagi, 2003). On the other, for inflation parameters, it is worth considering that in 1992 wage indexation was reformed in Italy and therefore a structural break between 1992 and 1993 is what is reasonable to expect. The reform of wage indexation determined the change from a more forward looking Phillips curve to a more backward looking one.

The model by Gali and Gertler (1999), though partially rejected by the data, could be useful to disentangle this puzzle. In their model γ_f and γ_b are a function of the structural parameters:

$$\gamma_f \equiv \frac{\beta\theta}{\theta + \omega\left[1 - \theta\left(1 - \beta\right)\right]}$$
$$\gamma_b \equiv \frac{\omega}{\theta + \omega\left[1 - \theta\left(1 - \beta\right)\right]}$$

where β is the discount factor, θ is the share of agents that in a Calvo price setting equation are locked in past contracts, ω is the share of backward looking price setters. Wage indexation in Italy before 1993 had a quarterly frequency and it was set by a commission. After 1993 it became the outcome of the bargaining process between trade unions and firms. It is likely that, before 1993, the high frequency and the public nature of the process of the adjustments limited both the value of the past level of prices as predictors of future ones and the bargaining costs that unions and firms might incur when setting a new wage, reducing the share of backward lookers and therefore γ_b.[16] One further explanation for this structural break could rely on the marked decrease that the

[16] Furthermore, from this point of view Galì and Gertler (1999) model is preferable

average inflation rate experienced between 1992 and 1993. Indeed, the literature on output persistence have showed that the higher is trend inflation and the lower is persistence, because agents have an incentive not to stick to old contracts (Ascari, 2000). So the literature on indexing might not have thoroughly considered the effect of trend inflation and wage bargaining on economic agents' choices.

More stable results can be obtained by splitting the sample between North and South (Columns 8 and 9). In the "North" model unemployment loses significance. However, a series of Wald tests could not reject the null that inflation expectations, the inflation lag and the unemployment rate could be pooled (for inflation expectations it returned a value of 0.45 with a p-value of 0.5, for inflation lag a value of 2.69 with a p-value of 0.10 and for unemployment a value of 0.18 with a p-value of 0.67). Moving the border regions like Tuscany, Umbria and Marche from the South subsample to the North subsample would not change the results much and a joint Wald test could not reject poolability, returning a value of 2.86 with a p-value of 0.41.

As showed in Column 10 of Table 2, I reduced the number of instruments, because one of the critiques moved to the New Keynesian literature is too heavy instrumentation. To check if too many instruments badly affect estimates, I used only two lags of inflation and results are stable in terms of size, sign and significance.

Finally, Column 11 in Table 2 shows the results for an AR(1) model of inflation, comparable for instance to the model estimated by Weber and Beck (2003). With difference to the previous specifications that can be regarded as conditional $\beta-$convergence equations, the specification in Column 11 is an absolute $\beta-$convergence equation. It is worth noting that the results in Column 8 are robust to spatial heterogeneity, due to the adoption of the system GMM estimator, and that the Arellano and Bond tests do not detect serial correlation in the original residuals, therefore the inclusion of time dummies here is unnecessary. Their insertion, however, would not change the results and estimates are available from the author upon request.

Comparing the coefficient estimates of Column 11 with those of Column 4, it appears clear that the danger of estimating absolute convergence equations instead of conditional ones is a substantial overestimation of inflation persistence. Estimating absolute convergence models may lead to think that inflation rates are converging slowly to the same value and in fact they are converging to different values in different

to the one by Christiano, Eichenbaum and Evans (2005), where the weights of the inflation lag and of inflation expectation depend only on the discount factor, which is unlikely to change from one year to the other.

places, conditional on the forcing variable. This result is not only against the purchasing power parity hypothesis, but also against the view of a long-term vertical Phillips curve. The long-term inflation rate will be different in different regions only if $\gamma_f + \gamma_b \neq 1$:

$$\hat{\pi}_i = \frac{\lambda}{1-\gamma_f-\gamma_b}\hat{x}_i + \frac{\mu_i}{1-\gamma_f-\gamma_b}$$

This could help to explain why long term processes of real appreciation or depreciation of regional economies take place, as highlighted by non-parametric results. Furthermore, this could shed more light on why models such as those in Cecchetti et al. (2002) or Weber and Beck (2003), omitting the role of the forcing variable, due to data constraints, and of future expectations, point to a large persistence of inflation or price differentials, while the descriptive statistics presented in the very same papers depict inflation as a volatile phenomenon. Finally, the insertion of the forcing variable, as unemployment, does not only help to have results more easily reconcilable with descriptive statistics, it also helps to reconcile better the presence of inflation β-convergence and the absence of a smooth inflation $\sigma-$convergence. Indeed, (6) does not only helps to go beyond just saying that inflation dispersion depends on the dispersion of the disturbance like in absolute convergence models, it also highlights that inflation dispersion depends on the dispersion of future inflation expectations, on the dispersion of the first lag, on the dispersion of the forcing variable and in the covariances of the regressors:

$$\begin{aligned}\sigma^2_{\pi t} = {} & \lambda^2\sigma^2_{xt} + \gamma_f^2\sigma^2_{E\pi t+1} + \gamma_b^2\sigma^2_{\pi t-1} + \sigma^2_{\epsilon t} + \\ & +2\gamma_f\gamma_b Cov(E\pi_{t+1}, \pi_{t-1}) + 2\lambda\gamma_b Cov(x, \pi_{t-1}) + 2\lambda\gamma_f Cov(x, E\pi_{t+1})\end{aligned} \tag{16}$$

where σ^2 is the symbol for the variance. Therefore, changes in the variance of expectations or in the variance of the forcing variable or even in their covariances may explain why we do not observe a smooth sigma convergence in the data. This implies a pessimistic message regarding inflation convergence across different regions of a monetary union as long as there exist regional disparities in terms of unemployment or expectations there will never be a complete convergence.

However, the heterogeneity of inflation expectations is known to depend on the level of inflation (Heymann and Leijonhufvud, 1995; Cukierman and Meltzer, 1986 and Ball, 1992) and this could explain why, like in Weber and Beck (2003), also analysing the dataset here proposed a positive correlation between the average and the dispersion of inflation emerges. By using Newey-West standard errors robust to heteroskedasticity and serial correlation it is possible to obtain the following result:

$$E\left(\sigma_{\pi t} | \mu_{\pi t}\right) = \underset{(0.01)}{0.33} + \underset{(0.00)}{0.11} \mu_{\pi t} \tag{17}$$

where $\sigma_{\pi t}$ is the standard deviation of inflation and $\mu_{\pi t}$ is the average inflation rate.

This result, though tentative because based on a small sample, is the same as that of Weber and Beck (2003) and it carries a great importance implying that central banks can reduce the real interest rate imbalances between high inflation regions and low inflation ones by decreasing the aggregate inflation rate and they can do it without a sizeable portion of regions within the monetary union to fall into deflation.

As showed by Weber and Beck (2003), using the equation for the dispersion of inflation and assuming that within each cross-section inflation is normally distributed[17], it is possible to compute the critical inflation rates that force a given percentage of regions into deflation. Table 3 shows the results for the present study. Though not adjusting inflation standard deviation, a 2% average inflation rate will entail that more than 2% of regions is in deflation, adjusting it according to (17), the 2% of regions will be in deflation with a 0.8% average inflation rate, due to the reduction in inflation regional dispersion. Therefore, though inflation rates may follow different paths in different regions, monetary policy can affect the dispersion of these paths.

7 Conclusions

The main purpose of this paper is to merge together two strands of the literature regarding inflation, the PPP and the Phillips curve ones, in the attempt to reassess some of their open questions. As a consequence a number of issues have been raised.

In the first place, in accordance with the finding of the NKPC literature and with the descriptive statistics usually provided in the PPP literature, but contrary to the findings of the regression analyses of the PPP literature, with the exception of Imbs et al. (2005), inflation persistence appears to be rather small: the implied half-life of the results of this contribution, computed as $\frac{\log 0.5}{\log \gamma_b}$, is 0.42 years or 5 months. However, inflation is not characterized by absolute β-convergence, but by conditional β-convergence, because as highlighted by the literature above there are not only "equilibrating" mechanisms, such as arbitrage, but

[17] The hypothesis of normal distribution was rejected only for the years 1987 and 1995 by a skewness-kurtosis test run for all the years of the sample. Therefore, it is possible to think that normal cross-sectional distribution can be considered a good approximation of the data also in this application.

also “dis-equilibrating” ones, such as different real interest rates in different regions.

Comparing different models proposed by the traditional and the New Keynesian Phillips curve literature, it is possible to see that in this application “old wisdom” beats “new knowledge”. In other terms, though non parametric analysis of the data would allow to think that the “cost based” or the “output gap” based Phillips curve could equally fit the data as the traditional Phillips curve, regression analysis shows that this is in fact superior because unemployment is the only significant forcing variable.

The comparison regarding what is the most suitable model specification was not only fruitful to compare different economic schools of thought, it also allowed to understand which of the three explanations proposed by the regional PPP literature better suits the data. The fact that unemployment prevailed over the output gap and the real unit labour cost means that inflation differentials are not driven by either different cost structures or different market imperfections, rather by differences in effective demand in different regions possibly due to sectorial specialization of each region and credit market imperfections. As a consequence, it is safely possible to state that inflation convergence appears in this application to be conditional on unemployment.

Therefore, as highlighted by the literature regarding inflation differentials within EMU, regions with higher unemployment rates will experience a persistent real depreciation (re-equilibrating mechanism), however this will lead them to pay a higher real interest rate than regions with less unemployment. What is more there is evidence that, due to market imperfections, lagging regions already pay an higher nominal interest rate than more advanced one. For instance, Banca d’Italia (2000) estimated the nominal interest rate charged by bank branches in the North West of Italy in March 2000 to be 5.17%, in the North East 5.77, in Central Italy 6, in the South 7.34 and in the Italian islands 7.30. This also entails that "equilibrating" and "dis-equilibrating" mechanisms pass through different markets: arbitrage mainly works through the goods market, whereas the real interest rate channel works mainly through the capital market.

Merging the Phillips curve and the PPP literature also allowed to reconsider the issue of the long run vertical Phillips curve. In the light of the recent contributions questioning the existence of a vertical long run unemployment Phillips curve, such as Mankiw (2001), Graham and Snower (2003) and Vaona and Snower (2006) and, especially for a regional setting, Hughes-Hallet (2000), I found evidence of a non-vertical unemployment-inflation relationship when considering a regional dataset. Therefore, demand redistribution between regions could help to reduce the long-run aggregate unemployment rate.

Furthermore the PPP literature overlooked the importance of future inflation expectations. This could explain the puzzles of mobile regional inflation rates converging slowly and the absence of a smooth sigma inflation convergence. These descriptive features cannot be found in unemployment and so, though this variable can help to explain why some regions may experience long lasting deflationary processes, it is unlikely to give a thorough explanation for the volatility and the dispersion of regional inflation rates. To this purpose the fact that inflation expectations significantly affect present inflation rates is more promising, as changes in expectations can explain the high volatility of inflation rates.

Finally, the importance of inflation expectations may help to explain why there is a positive correlation between the dispersion and the average level of inflation, allowing central banks to reduce inflation without a sizeable portion of regions falling into deflation.

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Figure. 1: Standard Deviation and Average Level of the Italian Provincial Inflation Rates from 1987 to 1998

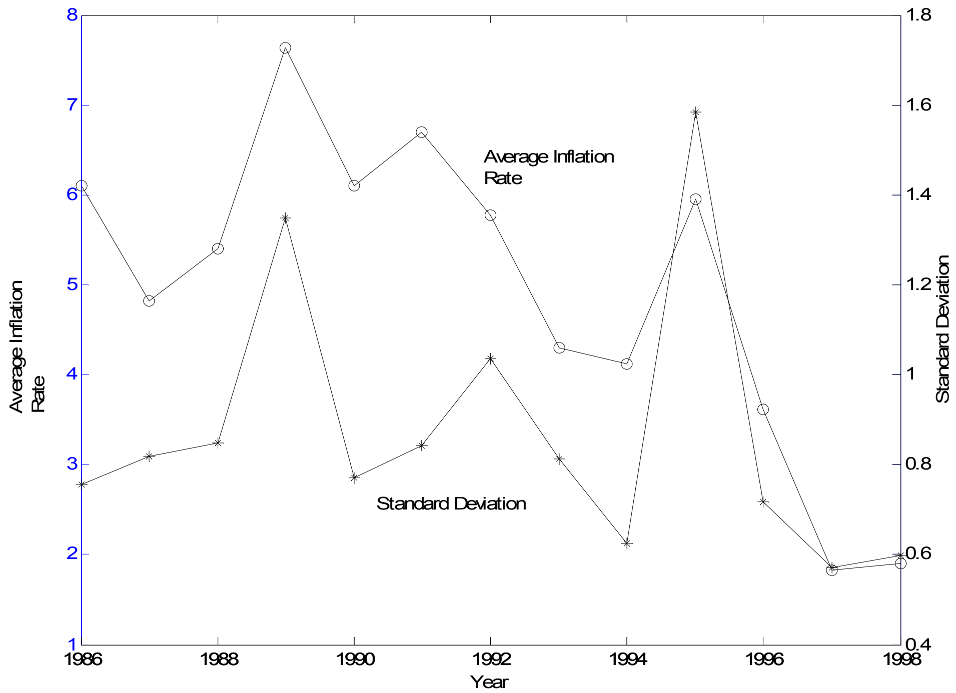

Figure 2: Inflation Rates Dynamics Across Italian Provinces: Estimated Stochastic Kernel (t-τ: 1987, t: 1998)

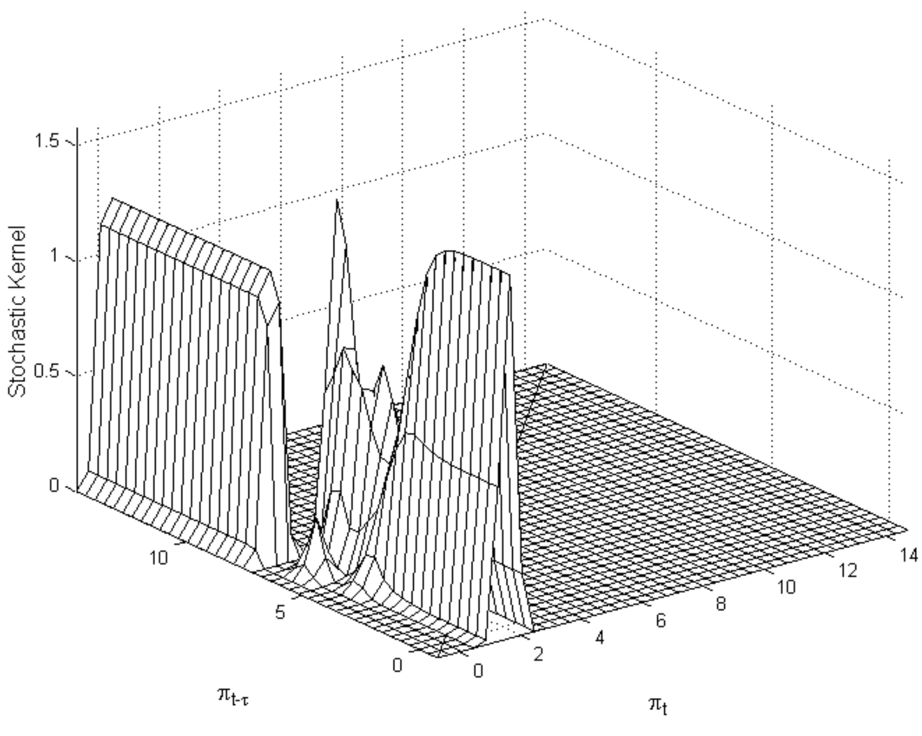

Figure 3: Inflation Rates Dynamics Across Italian Provinces

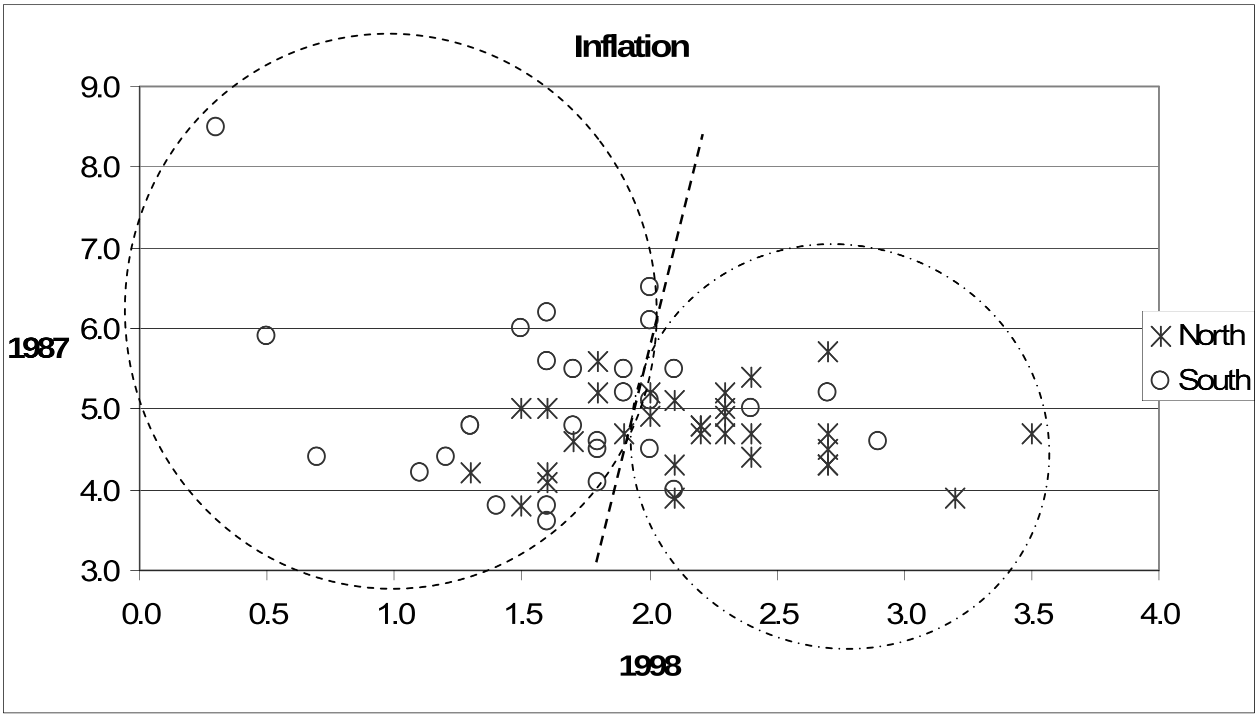

Figure 4: Value Added Dynamics Across Italian Provinces

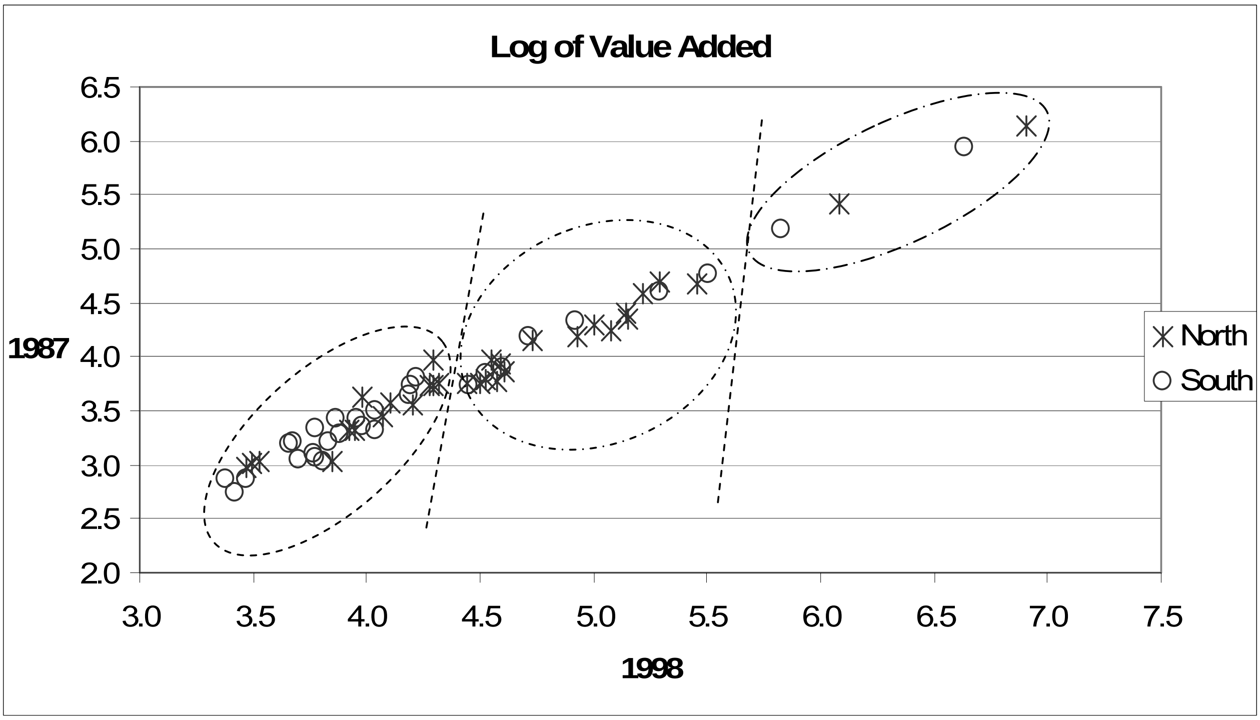

Figure 5: Dynamics of the Unemployment Rate Across Italian Provinces

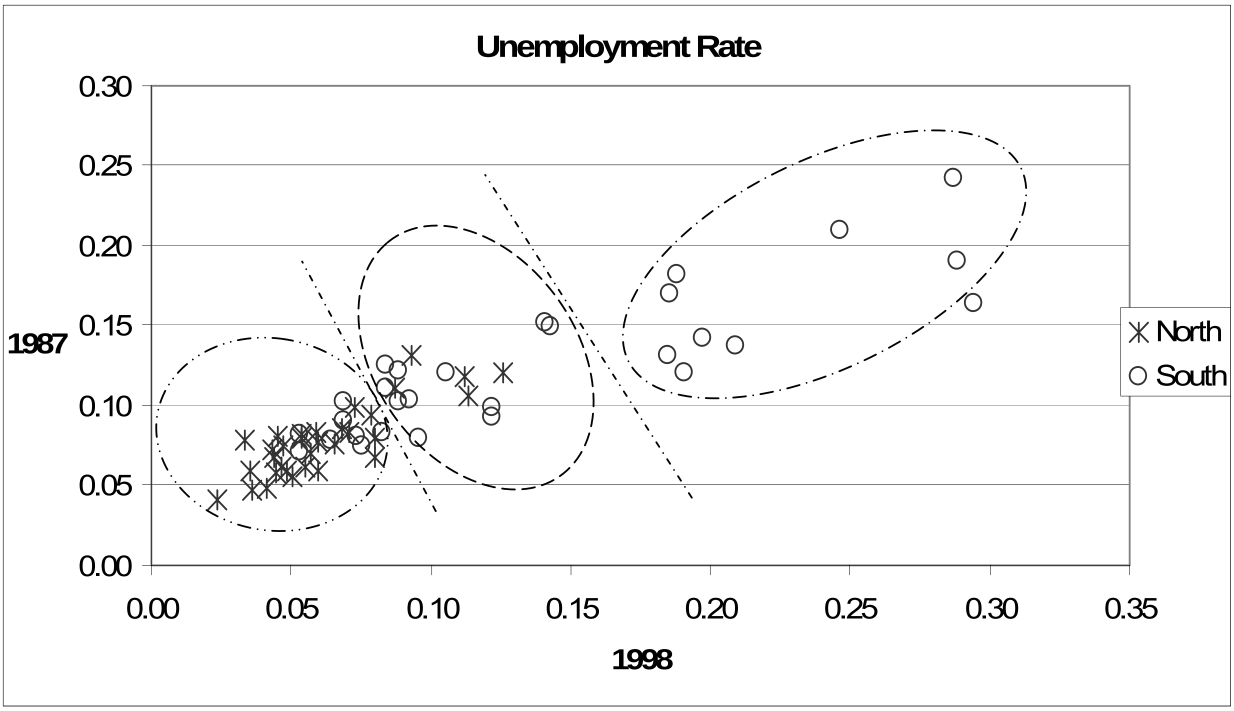

Figure 6: Dynamics of the Log of the Real Unit Labour Cost Across Italian Provinces

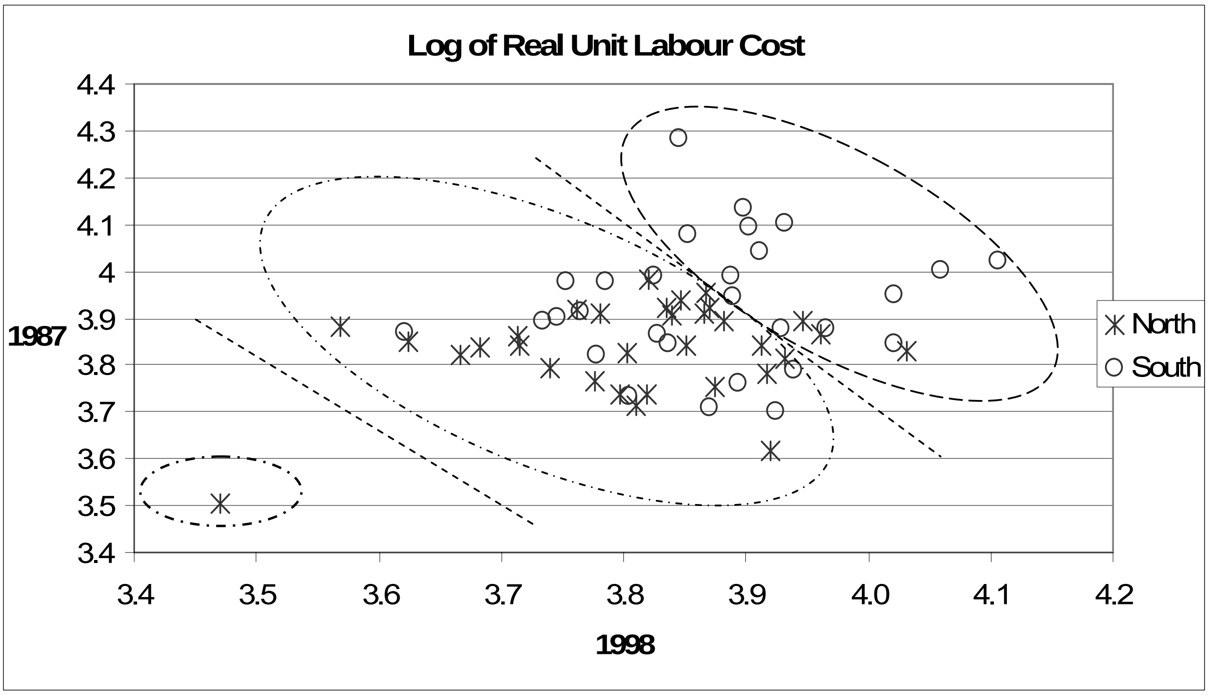

Figure 7: The Long-run Phillips Curve for Different Years

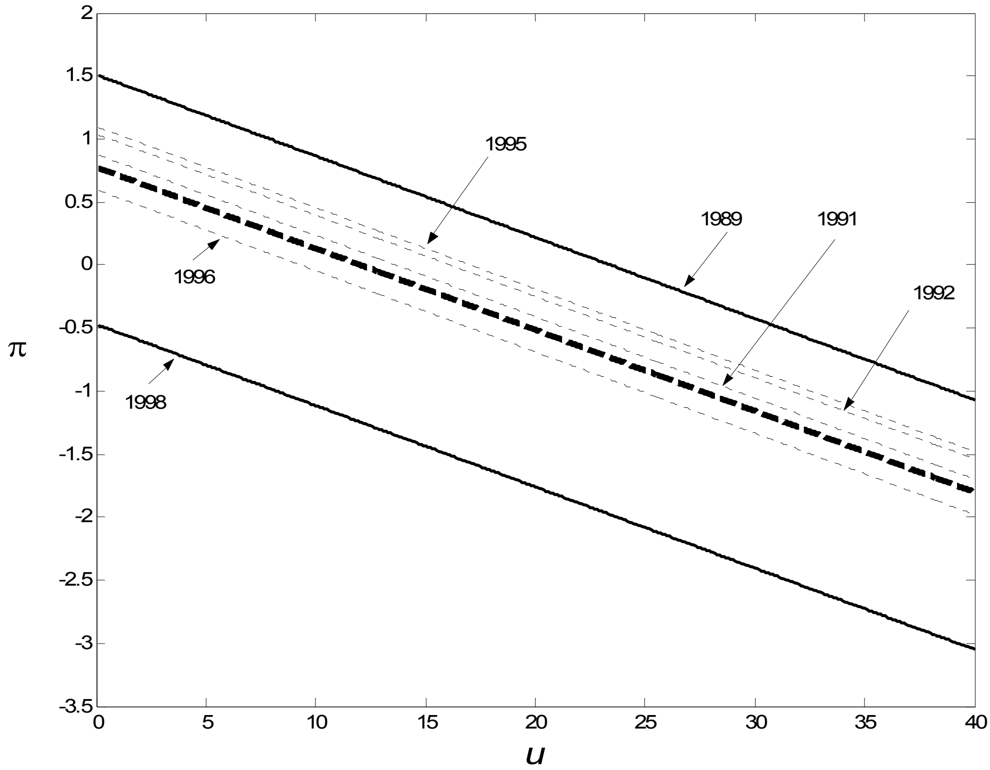

Table 1: Moran's I statistics for the inflation rate (π), the log of the value added (y), the unemployment rate (u) and the log of the real unit labour cost (mc) before and after spatial filtering (p-values in parentheses).

Variable	Before Filtering	After Filtering
π	15.37	-0.15
	(0.00)	(0.88)
y	15.24	-3.69
	(0.00)	(0.00)
u	9.39	-1.42
	(0.00)	(0.15)
mc	15.18	-2.44
	(0.00)	(0.02)

Table 2: System GMM estimates for different model specifications - Dependent variable: π_t (p-values in parentheses)

Variable	Model Specification										
	output	cost	unemp1	unemp2	unemp3	<1993	>=1993	North	South	few instr	AR(1)
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
π_{t+1}	0.50	0.51	0.48	0.50	0.52	0.60	0.41	0.43	0.57	0.53	-
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	-
π_{t-1}	0.14	0.15	0.15	0.19	0.18	0.19	0.29	0.13	0.22	0.23	0.69
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.02)*	*(0.00)*	*(0.00)*	*(0.00)*
y	0.06	-	-	-	-	-	-	-	-	-	-
	(0.10)	-	-	-	-	-	-	-	-	-	-
mc	-	0.01	-	-	-	-	-	-	-	-	-
	-	*(0.97)*	-	-	-	-	-	-	-	-	-
u	-	-	-0.02	-0.02	-0.02	-0.01	-0.02	-0.02	-0.01	-0.01	-
	-	-	*(0.02)*	*(0.00)*	*(0.01)*	*(0.05)*	*(0.02)*	*(0.17)*	*(0.03)*	*(0.01)*	-
d1987	0.40	0.37	0.36	-	-	-	-	0.04	0.09	-	-
	(0.40)	*(0.43)*	*(0.44)*	-	-	-	-	*(0.94)*	*(0.93)*	-	-
d1988	-0.15	-0.20	-0.15	-	-	-	-	-0.25	-0.67	-	-
	(0.74)	*(0.68)*	*(0.75)*	-	-	-	-	*(0.62)*	*(0.17)*	-	-
d1989	1.70	1.66	1.70	1.51	1.39	1.53	-	1.34	1.10	1.47	-
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.01)*	*(0.01)*	-	*(0.00)*	*(0.00)*	*(0.00)*	-
d1990	0.60	0.56	0.61	-	-	-	-	0.13	-0.02	0.87	-
	(0.18)	*(0.20)*	*(0.16)*	-	-	-	-	*(0.74)*	*(0.75)*	*(0.01)*	-
d1991	1.14	1.09	1.13	0.88	0.81	0.95	-	0.51	0.67	-	-
	(0.01)	*(0.01)*	*(0.01)*	*(0.00)*	*(0.00)*	*(0.01)*	-	*(0.05)*	*(0.02)*	-	-
d1992	1.31	1.29	1.30	1.04	1.03	1.26	-	0.52	0.77	1.03	-
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.01)*	-	*(0.02)*	*(0.06)*	*(0.00)*	-
d1993	0.42	0.40	0.40	-	-	-	-	0.25	-0.13	-	-
	(0.16)	*(0.18)*	*(0.17)*	-	-	-	-	*(0.46)*	*(0.55)*	-	-
d1994	0.11	0.09	0.11	-	-	-	-	0.41	-0.36	-	-
	(0.75)	*(0.78)*	*(0.73)*	-	-	-	-	*(0.24)*	*(0.18)*	-	-
d1995	1.24	1.23	1.25	1.10	1.02	-	0.88	0.23	1.44	1.62	-
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	-	*(0.00)*	*(0.32)*	*(0.00)*	*(0.00)*	-
d1996	0.80	0.80	0.80	0.60	0.59	-	-	0.48	0.27	-	-
	(0.03)	*(0.03)*	*(0.03)*	*(0.02)*	*(0.05)*	-	-	*(0.16)*	*(0.40)*	-	-
d1997	0.05	0.05	0.05	-	-	-	-0.45	0.09	-0.09	-	-
	(0.81)	*(0.81)*	*(0.81)*	-	-	-	*(0.00)*	*(0.66)*	*(0.62)*	-	-
cons	-0.67	-0.65	-0.67	-0.47	-0.43	-0.65	-0.13	-0.30	-0.30	-0.46	-
	(0.02)	*(0.02)*	*(0.02)*	*(0.00)*	*(0.00)*	*(0.02)*	*(0.05)*	*(0.07)*	*(0.13)*	*(0.00)*	-
Sargan Test	67.52	68.79	66.63	72.14	72.14	13.19	57.51	21.69	32.00	37.07	78.13
	(0.36)	*(0.32)*	*(0.39)*	*(0.41)*	*(0.41)*	*(0.58)*	*(0.31)*	*(1.00)*	*(1.00)*	*(0.07)*	*(0.41)*
AB test for AR (1)	-5.82	-5.83	-5.85	-6.04	-5.64	-3.64	-5.53	-4.73	-4.44	-5.75	-5.54
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*
AB test for AR (2)	1.78	1.80	1.71	1.77	1.56	1.75	1.91	2.77	0.97	1.39	1.51
	(0.08)	*(0.08)*	*(0.09)*	*(0.08)*	*(0.11)*	*(0.08)*	*(0.06)*	*(0.00)*	*(0.33)*	*(0.16)*	*(0.13)*
Moran's I	1.80	*1.80*	*1.83*	*1.83*	*1.67*	-	-	-	-	-	-
	(0.07)	*(0.07)*	*(0.07)*	*(0.07)*	*(0.10)*	-	-	-	-	-	-
Obs.	868	868	868	868	868	430	438	448	420	868	868
N. of groups	81	81	81	81	81	76	77	39	42	81	81
N. of instr	79	79	79	79	79	22	59	79	79	34	77

Table 3: Share of regions in deflation for different average inflation rates (assuming a cross-sectional normal distribution for inflation)

$\pi_{avg.}$	Adj. σ	Non adj. σ
2.0	0.0001	0.0228
1.9	0.0002	0.0287
1.8	0.0003	0.0359
1.7	0.0005	0.0446
1.6	0.0008	0.0548
1.5	0.0012	0.0668
1.4	0.0019	0.0808
1.3	0.0030	0.0968
1.2	0.0047	0.1151
1.1	0.0074	0.1357
1.0	0.0115	0.1587
0.9	0.0180	0.1841
0.8	0.0278	0.2119
0.7	0.0427	0.2420
0.6	0.0649	0.2743
0.5	0.0970	0.3085
0.4	0.1424	0.3446
0.3	0.2043	0.3821
0.2	0.2850	0.4207
0.1	0.3847	0.4602
0	0.5000	0.5000

Note: non-adjusted σ is equal to one.