# Inflation Dispersion and Convergence in Monetary and Economic Unions: Lessons for the ECB[†]

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## Abstract

Using a set of regional inflation rates we are examining the dynamics of inflation dispersion within the U.S., Japan and across U.S. and Canadian regions. We find that inflation dispersion is significant throughout the sample period. This findings suggests a "regional dimension" of the factors determining inflation. Based on methods applied in the empirical growth literature we provide evidence of significant mean reversion ($\beta$-convergence) in inflation rates in all considered samples. The evidence on $\sigma$-convergence is mixed, however. We argue that observed declines in dispersion are probably not the result of diminishing asymmetries across regions but are caused by decreasing overall inflation levels. This indicates a positive relationship between mean inflation and overall inflation dispersion. Our findings for within-distribution dynamics suggest that large deviations from mean inflation will decline with a probability of around 50%. Within-distribution dynamics is largest for Japanese prefectures, followed by U.S. metropolitan areas. For the U.S./Canadian sample, we find a degree of within-distribution dynamics comparable to that of EMU. We argue that this evidence of relatively smaller dyanamics in EMU does not necessarily indicate the inadequacy of a single monetary policy for EMU countries.

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# 1 Introduction

One of the major challenges that the European Central Bank (ECB) faces is the extent and dynamics of inflation dispersion in the Euro area. Critics of the establishment of a monetary union in Europe have expressed considerable doubts that a single monetary authority can adequately meet the requirements of such a heterogeneous group of countries as the member countries of the European Monetary Union (EMU) constitute ("Does One Size Fit All?"-debate). One issue that has been particularly been paid attention to in this context are the implications of Balassa-Samuelson effects for inflation dispersion. It is argued that - as a consequence of convergence in living conditions - inflation rates in "poorer" member countries such as Ireland or Portugal will be relatively high compared to "richer" member countries such as France or Germany. The resulting inflation dispersion constitutes a large problem for the ECB: When it strictly sticks to its target (an EMU-wide average inflation rate of less than 2%) several countries might face negative inflation rates. However, when it tolerates an EMU-wide average inflation rates of above 2% it looses credibility. In light of this dilemma Sinn and Reutter (2001) call for an increase in the ECB's upper target by at least 0.5%. Interestingly, the issue of inflation dispersion and convergence has played only a minor role in the literature before the establishment of EMU. One reason for this neglect might be that in well-established monetary unions such as the USA, Germany or France, regional inflation dispersion has not been present or of negligible size. That this is not true is clearly shown in Cecchetti et al. (2000) who find large and persistent differences in inflation rates across U.S. metropolitan areas throughout the last century with no trend to decline. As we will see later, this result is confirmed for all countries included in this work. Another reason for the neglect of regional inflation dispersion might be that all well-established monetary unions were taken as a datum and were more or less considered to be optimum currency areas. Thus, research in monetary policy focussed less on inflation dispersion within the monetary union but on other aspects such as strategic interactions between policy makers and the public, the right handling of uncertainties, ... The view of existing monetary unions as optimum currency areas is e.g. challenged by Rockoff (2000) who argues that the long-run existence of the U.S. monetary union is "at best weak evidence that the net effects (of the monetary union) have bee positive". Due to Rockoff (2000), only political - not economic - reasons have justified the maintenance of the U.S. monetary union.

As there is no example for the establishment of a monetary union of comparable size in recent history, assessments of current or potential future developments in the EMU area are difficult. One way to overcome this problem is to refer to evidence from well-established monetary unions. This might be helpful for several reasons both in general and in particular for the issue of inflation dispersion. In our opinion, the study of inflation dispersion in well-established monetary unions can generate important contributions to the ongoing debate on EMU inflation dispersion. First, the comparison can help us to assess the overall size of European-wide dispersion relatively to the U.S. regional inflation dispersion, e.g. This helps us to understand whether prevailing European distribution is outraging or of comparable (and thus sustainable?) size. Furthermore, the study of long existing (but not necessarily optimum) monetary unions can give us a hint where EMU inflation dispersion might evolve over time. Therefore, this paper considers dispersion evidence of two monetary unions (USA and Japan) and an economic union (USA and Canada as part of NAFTA) and tries to draw some lessons for the monetary policy of the ECB. Of course, due to significant differences between the considered samples and the EMU (such as a missing centralized fiscal authority), these comparisons have to be considered with some caution.

Both the theoretical and the empirical existing literature on regional inflation dispersion is very limited. Amongst the rare theoretical work on the sources and the dynamics of regional inflation differentials is Duarte and Wolman (2002). In their paper, the two authors show that inflation differentials in a monetary union can best be explained by productivity growth differentials across regions whereas fiscal policy does not seem to play an important role. Empirical work on U.S. evidence include Parsley and Wei (1996) and Cecchetti et al. (2000). Both of these studies use regional U.S. price data and show that shocks to relative prices are persistent. Cecchetti et al. (2000), e.g., find half lives of PPP deviations for U.S. cities of almost nine years which they consider to be a lower bound for EMU prices. Evidence for inflation dynamics within the Euro area is provided by Weber and Beck (2003) who use regional inflation rates from major EMU countries for the period before and immediate after the introduction of the Euro. The two authors show that there is considerable within distribution dynamics in European inflation rates. Furthermore, evidence is presented that inflation dispersion has reduced in the early 1990s and has reached a level compatible with the ECB's inflation target.

In this paper, we will contribute to this literature in several ways: First, we will examine the extent of mean-reverting behavior in regional inflation rates in the two well-established monetary unions USA and Japan and in the economic union between the U.S. and Canada. In a second step, we will examine the dynamics of overall dispersion in the our three samples for the last twenty years. Our main focus will be on the question how overall dispersion has evolved over time and how large dispersion in these samples is relative to the European evidence (documented in Weber and Beck (2003)). In our third contribution, we will focus on the question how the shape and the composition of regional inflation dispersion in our samples has evolved over time. To do so, we refer to a methodology known as distribution dynamics in the empirical growth literature. In our last contribution, we provide "critical values" for average inflation rates. These values serve to indicate when a significant portion of the involved regions face negative inflation rates. The results are then compared to the EMU case. All these questions will be examined using regional inflation data of the included countries that has not been used in the literature for these purposes before. The idea to use regional instead of national data is borrowed from the growth literature[1] where it has proven to be very useful in analyzing convergence in per-capita incomes.

The rest of the paper is organized as follows: In the next section, we present our data set and discuss some descriptive statistics. The results concerning mean-reverting behavior in inflation rates are presented in the following section. Section 4 examines the issue of $\sigma$-convergence in inflation rates and section 5 examines within-distribution dynamics. Section 6 takes a closer look at the relationship between the cross-sectional mean inflation rate and its dispersion and derives the above mentioned "critical values". The last section summarizes our results and draws some policy conclusions.

## 2 Data and Descriptive Statistics

As just mentioned, one important aspect of our analysis is the use of regional instead of national data. The regional perspective enables us to address questions that could not or not adequately dealt with on the basis of national data. The merits of such an approach can be readily seen from its use in the empirical growth literature where it has proven to provide important insights into the nature of per-capita income across regions[2]. But also in international economics, this approach has proven to be useful. Authors like Engel and Rogers (1996), Parsley and Wei (1996) or Beck and Weber (2001) use it to examine questions concerning the integration of national and international goods markets. There are several reasons that make this approach so appealing to researchers: First, the use of regional data increases the number of observations and thus provides us with more precise statistical results. Secondly, the extra - regional - data dimension can help us to address questions that couldn't be dealt with otherwise. A comparison between within-country and cross-country goods market integration, e.g., as done in Engel and Rogers (1996) can only be done with regional data from a country. Likewise, an investigation of inflation dispersion in EMU as done in Weber and Beck (2003) cannot be seriously taken into account on

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[1] See e.g. Barro and Sala-i Martin (1992), Barro and Sala-i Martin (1995) and Sala-i Martin (1996a).

[2] See e.g. Barro and Sala-i Martin (1992), Barro and Sala-i Martin (1995) and Sala-i Martin (1996a) amongst others for a reference.

the basis of only 12 national CPI observations. The third reason why we use regional inflation data is the following: As each country can be considered as a miniature monetary union, the use of regional data from well-established monetary unions can probably give us some insights in further developments within EMU. Here, the study of U.S. cities is probably most helpful. A drawback of this approach is that regional data are not readily available and thus have to be collected in a time-consuming (and sometimes costly) process. Furthermore, even if one is willing to carry this burden, one may not be successful because some countries' statistical offices do not compile data at a regional level. Fortunately, the latter is not true for the countries included in our study. As we will see, the statistical offices of all included countries not only compile regional data but do so in a satisfyingly broadness.

We have compiled regional inflation data for the United States, Canada and Japan. An overview of included regions, the respective sample periods and the data sources is given in table 1. As one can see there, our data comprise 24 metropolitan areas in the U.S., 12 provinces in Canada and 47 prefectures from Japan. U.S. and Canadian CPI data are available for the time period 1980 - 2001, Japanese data are available for the period 1985 - 2001 [3]. In our estimation analysis, we are looking at three different samples constructed from these data: the U.S., Japan and the U.S. jointly with Canada. The first two are examples of monetary unions whereas the third sub-sample represents an economic union between two countries. These two countries do not share a common monetary policy but are closely linked both by economic and cultural linkages. The study of U.S. and Japanese regional inflation dynamics is motivated by two major aspects: First, it allows us to set results found in Weber and Beck (2003) for EMU in relation to the evidence from well-established monetary unions where within-country inflation dispersion is not seen as a problem. Secondly, it provides us with a benchmark of what we might expect for the EMU in the medium run future. The joint U.S./Canadian sample is chosen for the following reason: Due to the lack of a common monetary policy, this sample can serve to a certain extent as providing an antipole to the likewise considered monetary unions. We expect EMU dispersion lying somewhere in between these two poles where we would not be surprised when European evidence would be closer to the U.S./Canadian sample.

All data are annually and are available in index form. Inflation rates $\pi_t$ are computed as annual percentage changes in the price index as:

$$\pi_t = 100 * (\ln P_t - \ln P_{t-1}), \tag{1}$$

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[3] See the notes in table 1 for some exceptions.

where $\pi_t$ denotes the inflation rate in period $t$, and $P_t$ represents the respective index in $t$.

To get an idea of the extent and evolution of regional inflation dispersion, figures 1 to 3 plot inflation rates for all three samples. As is clear from these plots, regional inflation dispersion is not only of significant size but also is fairly persistent. For the U.S., e.g., the difference between the lowest and the highest local inflation rate is about 3%. For Japanese prefectures, the comparable number is somewhat smaller, but is still of significant size (about 1.5%). Not surprisingly, differences are largest for the U.S./Canadian sample. Comparing plots, one can see that the three figures share a similar pattern. Inflation rates are highest at the beginning of the 1980s and drop down until the mid 1980s. The subsequent increase until 1990/1991 is followed by a long-lasting smooth reduction until the second half of the 1990s. In recent years, there has been a slight increase in inflation rates. Looking at the "bandwidth" of reported inflation rates, we can observe an important difference between the North-American samples and the Japanese sample: For the U.S. and the U.S./Canadian sample, there are some indications for a decrease in overall dispersion. This decrease is particularly evident for the first years of the sample and is related to the significant reduction in average inflation rates at that time. After 1985, there are no longer signs of a declining overall dispersion. We will discuss this issue in more detail in section 6 where we will argue that the decline of dispersion at the beginning of the 1980s is probably a consequence of the decrease in the overall inflation level. One issue cannot be addressed by pure inspection of fiugres 1 to 3 is whether the composition of dispersion is changing over time. Of particular interest for us is the question whether the lower and upper tails of the distribution always contain the same regions. This question is important for policy makers as this would imply persistent or even permanent divergence in price levels across regions.

Table 2 reports period-average inflation rates and its respective cross-sectional dispersion. Looking at the 5-year subperiods, it becomes clear that mean inflation has continuously declined for the U.S. and the U.S./Canadian sample (from an average of about 5.5% to an average of about 2.5%) in the last twenty years. The inflation pattern for Japan is relatively different. As one can easily see, overall inflation has always been far below that of the U.S. and Canada with an maximum average inflation rate of 1.26% for the 1991 to 1995 period and an extremely low average inflation rate of only 0.30% in the last subperiod. Comparing reported dispersion measures, we can see that - not surprisingly - dispersion is highest for the U.S./Canadian sample and lowest for the Japanese data. An interesting feature of the reported data is that dispersion is always considerably lower when longer time periods are considered. This observation is in line with Cecchetti et al. (2000) and gives us a first hint with respect to the nature of within-distribution dynamics of regional inflation rates: When taking a long run perspective, inflation differential across regions seem to reduce. This is even true for countries that are linked only economically but not by a single monetary policy. However, at a short and medium run horizon, there is significant dispersion. The main task of this paper is to shed some light on the nature of the dynamics in inflation dispersion across regions.

Before, however, we briefly want to address the question of possible reasons for the observed inflation dispersions. To do so, we refer to the factors that the literature on inflation considers to be responsible for observed inflation and ask whether these factors have a "regional" dimension in the sense that they can be made responsible for inflation differences across individual regions. As a side effect we get an idea of how persistent existing inflation differentials can be expected to be. The persistence is directly related to the length of time that underlying factors are assumed to be effective. In figure 4 we give an overview of possible inflation determinants[4]. We grouped individual factors into four categories, namely demand-side factors, supply-side factors, institutional, political and cultural factors and expectations and market frictions. As the arrows indicate, we believe that there are complex feedback mechanisms at work and that there are probably no truly exogenous factors. As an example for such a feedback mechanism, one can take the relationship between actual and expected inflation. Expected inflation has an influence on actual inflation (by reducing the demand for money, e.g.), but of course actual inflation also leads to adjustments of inflation expectations.

As we already discussed, the individual factors might differ with respect to the "level" at which they become effective. Money supply, e.g., is centralized in a monetary union and is uniformly determined for all regions. Wage determination on the other hand can have a regional effect and fiscal policy is effective on a national and regional level. Inflation dispersion will only arise when individual factors have asymmetric impacts across regions. To answer the question of how persistent these variations are we have to answer the question of how long the underlying factors are effective. As McCallum (1990) points out, monetary growth is the most important long run determinant of inflation rates whereas most other factors have only short or medium run impacts. So, as most of the factors that might have an impact on inflation have a "regional" dimension, in the short and medium run we should expect inflation differences. In the long run, however, existing differentials - at least in a monetary union - should vanish. In other words, we are expecting to find convergence of inflation rates in our empirical work, the speed of which depends on the extent of market rigidities and the persistence of asymmetries in the determining factors, however.

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[4] Detailed surveys on theories of inflation can be found in Laidler and Parkin (1975), Humphrey (1980) and McCallum (1990).

## 3 $\beta$-Convergence of Regional Inflation Rates

Figures 1 to 3 showed that there is considerable dispersion across regional inflation rates in the considered samples. This dispersion is only troublesome, however, when regions with relatively low/high inflation rates will stay in this position for an extended period of time. This would lead to an increasing divergence in regional price levels such that monetary authorities would face contradicting claims concerning their necessary response. In this paper, we follow two strategies that can give us an answer to that question: First, we examine mean-reverting behavior ($\beta$-convergence) using standard panel unit root methods and second, we use distribution dynamics (see section 5). Following the soccer-league picture by Sala-i Martin (1996a), studying the presence of $\beta$-convergence examines the question of whether (and how fast) the "inflation rank" of a region is changing over time. The extent to which this happens is determined by the persistence of the factors responsible for differences in inflation rates. As some of these forces, as e.g. indexing differences or productivity differentials, can be long lasting we should not be surprised to find relatively large persistence in inflation differentials.

Figures 5 to 10 illustrate our approach. There we plot average annual changes in inflation rates over the respectively considered sample period against initial inflation rates. When there is significant mean reversion in inflation rates, we should find a negative relationship. This would imply that higher initial inflation rates would be accompanied by (relatively) lower subsequent changes. For illustrational purposes, we also included an auxiliary regression line in each plot that represents fitted values from an OLS-regression of inflation changes on initial inflation rates. A first glance at all pictures - including the U.S./Canadian case - shows that there are clear indications of $\beta$-convergence in our samples. However, it is not clear whether there are differences in convergence speeds across samples. The plots for the sub-periods indicate that the pattern of mean reversion has been relatively stable over the last 15 to 20 years.

To address the issue of $\beta$-convergence more formally we refer to panel unit root methods as developed by Levin and Lin (1993)[5]. Given our sample of inflation rates $\pi_{i,t}$ (with $i = 1, 2, \ldots, N$ denoting the regions of our sample and $t = 1, 2, \ldots, T$ representing the time index), the test for inflation convergence is based on the following equation

$$\Delta\pi_{i,t} = \rho\pi_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j}\Delta\pi_{i,t-j} + \epsilon_{i,t}. \tag{2}$$

Here, $\Delta$ denotes the one-period change of a variable and $\theta_t$ represents a common time effect. $\epsilon_{i,t}$ is assumed to be a (possibly serially correlated) stationary idiosyncratic

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[5] For a more detailed description of the estimation and simulation process, see the appendix.

shock. The inclusion of lagged differences in the equation serves to control for serial correlation. As the subindex of $k$ indicates, we allow the number of lagged differences to vary across individuals, whereby the respective number is determined using the top-down approach suggested by Campbell and Perron (1991). To take control of cross-sectional dependence, we subtract the cross-sectional mean such that equation (2) becomes

$$\Delta\tilde{\pi}_{i,t} = \rho\tilde{\pi}_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j}\Delta\tilde{\pi}_{i,t-j} + \epsilon_{i,t}. \qquad (3)$$

$\tilde{\pi}_{i,t}$ denotes the deviation of region's $i$ inflation rate from the cross-sectional mean in period $t$ and is computed as

$$\tilde{\pi}_{i,t} = \pi_{i,t} - \frac{1}{N}\sum_{j=1}^{N} \pi_{j,t}. \qquad (4)$$

To see whether mean-reverting behavior in inflation rates is present, we test the null hypothesis that all $\rho_i$s are equal to zero against the alternative hypothesis that they are all smaller than zero. If we can reject the null hypothesis of nonstationarity, inflation rates exhibit mean reverting behavior. In this case, any shock that causes deviations from equilibrium eventually dies out. The speed at which this occurs, can be derived from the estimated value for $\rho$ (denoted $\hat{\rho}$). Given $\hat{\rho}$, half-lives of convergence are computed using

$$t_{half} = \frac{\ln(0.5)}{ln(\hat{\rho})}.$$

Unfortunately, for finite samples the estimates for $\rho$ are biased downward (see Nickel (1981)). To correct for this downward bias, we are using the adjustment factor Nickel is suggesting. Critical values for the test statistics are obtained using a parametric bootstrap based on 2000 simulations of the data-generating process under the null hypothesis.

Results are presented in table 3. There we report estimated half-lives both for the total period and 10-years subperiods. As one can readily see there is very strong evidence of mean reverting behavior in inflation rates in all considered samples. Ranking individual samples, we get the not surprising result that inflation convergence is largest in Japan, followed by the USA and the USA/Canadian sample. For Japan, we obtain half-lives of inflation deviations from the cross-sectional mean of around 6 months. For the U.S., half-lives lie slightly above and are in the range of 9 months. Interestingly, our results for the sub-periods indicate that half-lives might have considerably increased in the last decade. For the U.S./Canadian sample, half-lives are above 1 year. For subperiods, half lives are almost 2 years. As estimates for $\rho$indicate, the difference between results for the total period and subperiods are essentially due to the adjustment factor. As we found in our study of EMU inflation rates and is also documented in Cecchetti et al. (2000), the use of the adjustment factor for relatively short sample periods likely overestimates half-lives. We thus think that for the U.S./Canadian case, half-lives of slightly above 1 year are the more plausible case. Interestingly, differences in estimated half-lives between the U.S./Canadian sample and the two considered monetary unions are surprisingly small. To a large degree, this is probably an implication of the fact that Canadian monetary policy closely follows the U.S. monetary policy. Additionally, the already mentioned close cultural and economic linkages between these two countries make large asymmetries in inflation dynamics less likely.

Summarizing, our results for $\beta$-convergence indicate half-lives of deviations from the cross-sectional mean less than one year in well-established monetary unions and slightly more than one year for the considered U.S./Canadian economic union. Results for subperiods suggest that there is no trend towards a further decrease in half-lives. This can probably be seen as a sign that some type of steady state of inflation dispersion has been reached in the considered samples. Half-lives for the monetary unions show that there is some persistence in deviations of inflation rates from the cross-sectional mean, however, the degree of persistence is not troublesome for monetary authorities. This is especially true for Japan.

Comparing the reported figures with those found for EMU (see Weber and Beck (2003)), we see significant differences in the speed of mean reversion: Not surprisingly, half-lives in the Euro area are considerably higher than those for the two monetary unions that we consider in this paper. More surprisingly, however, is the fact that EMU half-lives are also significantly higher than those for the U.S./Canadian sample. Whereas differences to the U.S. and Japan could easily be ascribed to factors like the lack of a central fiscal authority, explaining the large differences to the U.S./Canadian case are more difficult. Depending on one's point of view these differences can be interpreted in several ways: For a critic of the Euro, it reflects the troublesome heterogeneity of countries in EMU and might be a sign that the monetary union was not a good idea. This line of argument is supported by the finding that mean reversion in inflation rates of two countries that do neither share a common monetary nor fiscal policy is higher than for EMU countries. A proponent of the Euro will assess the evidence differently. First, he will point out that there is inflation convergence in EMU, only the speed is relatively small. Secondly, he will - as we did above - argue that linkages between the U.S. and Canada are very close. Thus, the degree of heterogeneity across these countries might be of moderate size. Then, existing inflation dispersions across the U.S. and Canada do not necessarily represent unsustainable levels of dispersion. And last but not least, he will mention that EMU is still in a transition phase and has not reached a steady state, yet. In this sense, particularly results for the U.S. represent some benchmark toward which the Euro area will move in the long run. Unfortunately, we do not have long enough data to examine the latter view.

## 4 $\sigma$-Convergence of Inflation Rates

In the previous section, we examined the mean reverting behavior of regional inflation rates, i.e. we performed tests of $\beta$-convergence. In the growth literature, another concept of convergence, denoted as $\sigma$-convergence, has been extensively studied[6]. This concept focuses on the evolution of overall dispersion of the variable of interest. Applied to our case, this means to examine whether overall regional inflation dispersion has decreased, has remained constant or has increased during our sample period. For a central banker, the question of inflation dispersion is of great importance. When overall dispersion is high, he faces the difficult task to meet conflicting demands of different regions: Whereas for low-inflation areas an expanding monetary policy might be adequate, for high-inflation areas a restrictive policy is necessary. In this section, we will provide evidence on the dynamics of overall inflation dispersion for the U.S., Japan and the U.S./Canadian sample. Additonally, we will relate the findings to previously obtained results for the Euro area.

Following the empirical growth literature, we examine $\sigma$-convergence by computing the standard deviation of cross-regional inflation dispersion in each sample period. The results are plotted in figures 11 to 13. As one can easily see, the U.S. and U.S./Canadian samples on the one side and the Japanese sample on the other side considerably differ with respect to two important issues. First, the overall level of dispersion in Japan is around 0.4 and thus considerably smaller than that of the U.S. (0.6 and higher) or the U.S./Canada sample (0.8 or higher). This is true for the full sample length but particularly pronounced for the first sample years. Both the fact that these differences in dispersion exist and their size are not surprising given the different sizes of included countries. As the literature on national and international goods market integration shows[7], markets are regionally segmented with integration positively depending on the distance between markets. In light of these findings, it is not surprising that dispersion is more pronounced for the U.S. than for Japan. The second difference between the North-American samples and Japan concerns the dynamics of overall dispersion. For North-American samples, we find signs of decreasing overall dispersion. This is particularly evident for the deflation phase at

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[6] The terms $\beta$- and $\sigma$-convergence date back to the Ph.D. thesis of Sala-I-Martin, see Sala-i Martin (1990).

[7] see e.g. Engel and Rogers (1996) and Parsley and Wei (1996) for North-America and Beck and Weber (2001) for Europe.

the beginning of the 1980s, but even after this period there is still a further, however much smoother decline until the mid 1990s. From then on, overall dispersion slightly has increased. For Japanese prefectures, overall dispersion does not show a comparable decline. On the contrary, it has even slightly increased in the sample period. The interesting coincidence of movements in mean inflation and dispersion will be explored in more detail in section 6. Unlike in the case of $\beta$-convergence, the differences in overall dispersion between the U.S. (U.S./Canadian) sample and the EMU are not so pronounced. Weber and Beck (2003) show in their paper that there has been a significant dispersion in EMU regional inflation rate dispersion from about 1.5 to about 0.7/0.8. These number compare to 0.7 for the U.S. and 0.8 for the U.S./Canadian sample. Thus, overall dispersion in the Euro area is in the same range as it is for the well-established monetary union USA. This result is supportive to the upper mentioned Euro proponents' view. The smaller degree of mean reversion then may be the result of the well-documented larger rigidities in European markets.

Before proceeding, we want to clarify one issue. It concerns the maybe puzzling coexistence of strong evidence of $\beta$-convergence on one side and missing or "negative" evidence of $\sigma$-convergence for the U.S. or U.S./Canadian case (in the last few years) and Japan (throughout the sample period) on the other side. To do so, we take a closer look at the relationship between $\beta$- and $\sigma$-convergence. This is best be done by referring to Sala-i Martin (1996b) who shows that "$\beta$-convergence is a necessary condition for $\sigma$-convergence" but is not a sufficient condition for it. Following Sala-i Martin (1996a), the relationship between the two convergence concepts can be illustrated as is done in figure 14 that represents the different possible combinations of $\beta$- and $\sigma$-convergence. In panel a of figure 14, $\beta$-convergence induces $\sigma$-convergence. On the other hand, as panel b shows, the absence of $\beta$-convergence implies the absence of $\sigma$-convergence. The most interesting case is demonstrated in panel c. Here, $\beta$-convergence occurs, however, $\sigma$-convergence cannot be observed. In other words, the strong evidence of $\beta$-convergence found in the last section does not imply $\sigma$-convergence. In fact, it is - as we have seen - even possible to find increasing dispersion when "leapfrogging" occurs to a large extent.

## 5 Distribution Dynamics

In the last two sections, we found that overall inflation distribution has not reduced (or has even slightly increased) in the last few years, but that there are strong indications for mean-reverting behavior in inflation rates. These results imply that there is considerable within-distribution dynamics in regional inflation rates. In this section, we want to examine the nature of this within-distribution dynamics. To do so, we are using a methodology that is leant from the empirical growth literature where is has become known as *distribution dynamics*. One primary objective of its use in the growth literature was to study the composition of worldwide income distribution over time[8]. Analogously, its use in this paper serves to examine the evolution of the composition of regional inflation dispersions. Knowing how the composition of the tails of existing inflation dispersions changes over time is important for U.S. and Japanese policy makers for the reasons outlined above. Additionally, the results of this section can be useful for decision makers of the ECB. In Weber and Beck (2003), we perform a similar exercise using European regional inflation rates. We find that there is significant within-distribution: Large deviations from cross-regional means are expected to decrease considerably at a one-year horizon. Transition probabilities of staying at the left or right tail of the overall distribution are estimated to be about 50%. Comparing these results for EMU with analogous findings for the two monetary unions considered in this paper, can help to better assess the degree of dynamics in Europe. The U.S./Canadian case is - as outlined above - more considered to constitute an extremum that serves as an upper benchmark for EMU. Given the somewhat contradictory results of the last two sections concerning the assessment of EMU inflation dispersion, evidence in this section hopefully helps to clarify issues. The idea behind distribution dynamics is to find a law of motion that describes the evolution of the entire considered distribution over time. Following the growth literature, we are using a Markov processes to describe the dynamics of the cross-regional inflation distribution $F_t$ and model the dynamics of the cross-regional inflation distribution as an AR(1) process in the following way[9]:

$$F_{t+1} = T^{\star}\left(F_t\right). \tag{5}$$

Here, $T^{\star}(.)$ denotes the operator mapping period's $t$ distribution into period's $t+1$ distribution. Depending on the nature of the underlying variable of interest $X_t$, this operator is either interpreted as the transition function/stocastic kernel of a continuous state-space Markov process or the transition probability matrix of a discrete state-space Markov process. In the former case, equation (5) translates to

$$F_{t+1} = \int\limits_{A} P(x, A) F_t(dy), \tag{6}$$

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[8] See Bianchi (1997), Hobijn and Franses (2001), Quah (1993a), Quah (1993b), Quah (1996), Quah (1994) or Quah (1997) amongst others. For a recent survey, see Durlauf and Quah (1999).

[9] The following exposition is a condensed representation of the methodology of distribution dynamics. A more technical exposition can be found in Quah (1997) or in the appendix of Durlauf and Quah (1999).

where $A$ is any subset of the underlying state space for $X_t$ and $P(x, A)$ denotes the stochastic kernel that describes the probability that we will be in $A$ in $t + 1$ given that we are currently in state $x$, i.e.

$$P(x, A) = P(X_{t+1} \in A | X_t = x). \tag{7}$$

The variable of interest $X_t$ in our case is defined to be the deviation of a region's inflation rate from the cross-regional mean, the underlying state space is then the real line R.

We also consider a discretized case. A discrete-case consideration has the advantage that it provides us with easily interpretable transition probability matrices. The major drawback of this approach is, however, that any discretization will be somewhat arbitrarily. Insofar, the figures presented in this section have to be treated with caveat but are very useful for practical considerations[10]. For a discrete state-space, equation (5) becomes

$$F_{t+1} = MF_t. \tag{8}$$

$M$ is an $nxn$ transition probability matrix with $n$ denoting the number of distinct states and rows entries summing up to 1. Matrix entry $M_{ij}$ denotes the probability of a region's inflation rate that is currently in inflation state $i$ to move to state $j$ next period.

Figures 15 to 20 present the results for the continuous case. For each sample, the three-dimensional graph plots the the stochastic kernel for annual inflation rate transitions. On the x-axis (denoted by $t$), we plot period's $t$ inflation deviations from the cross-regional mean and on the y-axis (denoted by $t + 1$), we plot period's $t + 1$ inflation deviations from the cross-regional mean. On the z-axis, we plot the transition density function $p(x, y)$ associated with the stochastic kernel $P(x, A)$[11]. If the probability mass was concentrated along the diagonal of the x-y plain, then any existing deviations from the cross-regional inflation mean in period $t$ would be expected to remain basically unchanged over time. If on the other hand most of the probability mass in the graph was concentrated around the 0-value of the period-$t + 1$-axis - extending parallel to the period-$t$-axis - then period-$t$ deviations would

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[10] Another problem of discretization is that it can remove the Markov property (see e.g. Guihenneuc-Jouyaux and Robert (1998)). The results of Bulli (2000), that tries to evaluate the practical consequences of arbitrary discretizations, show that a regenerative discretization instead of our "naive" discretization would probably not change dramatically our main results.

[11] $p(x, y)$ has the property that

$$P(x, A) = \int_A p(x, y)dy, \tag{9}$$

with $y$ denoting elemetns in $A$. When $A$ is identical to the underlying state space (R), the transition density function integrates to one be basically expected to vanish until next period. The lines of the contour plots (left panel of the figure below the surface plots) represent lines of constant density of the respective surface plots. Conditional expected inflation rates for the period following the observation period are plotted to the right of the contour plots.

Looking at figures 15 to 20, we obtain a pattern of within-distribution dynamics that we is comparable to that from the section on $\beta$-convergence. However, differences between samples are more pronounced. The surface plot for U.S. metropolitan areas (figure 15) shows that the density mass is notably rotated clockwise. Thus, the conditional probability of a region with a large deviation from cross-regional inflation to stay in this position is relatively small whereas there is a large likelihood that it will be closer to the cross-regional mean in the next period. This can be seen very clear from the plot of next period's conditional expected inflation rate (right panel of figure 16). The conditional next period's expected mean deviation of a region whose mean deviation is $-2\%$ in this period is only $-0.5\%$. Thus, a considerable reduction in the deviation is expected. However, as the contour plot (left panel of figure 16) shows, there is also a non-neglectable probability that an existing deviation will not reduce. For Japan, the evidence of large dynamics towards the cross-regional mean is very pronounced. The surface plot (figure 17) is rotated by almost 45%. This means that any deviation from the cross-regional mean in the current period is more or less expected to vanish in the next period. This impression is confirmed by the plot of next period's condition mean deviation (right panel of figure 18). The contour plot (left panel of figure 18) shows that the probability of a large deviation from the cross-regional mean in period $t$ to persist until next period is very low. These patterns for within-distribution dynamics of regional U.S. and Japanese inflation rates are in line with our estimates of mean reversion. As we have argued there, the more pronounced dynamics towards the cross-regional mean in Japanese data is probably due to the smaller size - and thus the smaller degree of geographic segmentation - of this country. Looking at the results for the joint U.S./Canadian sample, results also confirm the smaller degree of dynamics towards the mean found in the section on $\beta$-convergence. Although, current deviations from the cross-regional mean are also expected to deviate, this is expected to happen to a significant less degree than for the U.S. This result is certainly not surprising as the regions of the two countries do not have a common monetary policy. An interesting conclusion is obtained when one compares the results for our three samples to those that we obtained in Weber and Beck (2003) for the EMU area. The best "match" is again found for the U.S./Canadian sample. The contour plots and the estimates for conditional expected changes in mean deviations between these two samples are similar. Thus, the monetary union EMU behaves more than the economic union U.S./Canada than it behaves relative to the monetary union USA. As we already argued above, this result probably reflects two features of the Euro area that have been often used in the past by Euro critics. The first is the absence of a central fiscal authority that can implement regional transfers to offset regional shocks. The second feature is the existence of very large market rigidities that additionally hinder the fast offsetting of asymmetric shocks. Together, these two features lead to more persistent effects of asymmetric shocks that is reflected in the weaker dynamics towards mean reversion in our inflation data.

While an examination of the continuous state space is more appropriate for inflation rates, a discretization has the advantage of providing figures for transition probabilities across states. These figures are especially useful for policy discussions. Thus, in the following, we provide transition probabilities for our three samples that we obtained by dividing the continuous state space into five discrete spaces. Boundaries for the spaces were chosen separately for each sample in a way that ensures that each state has an almost equal number of observations. Results are presented in table 4. The reported figures underpin the conclusion that we have drawn from the graphical findings for the continuous case. For the U.S., the probability of a region's deviation from the cross-regional mean to stay above 0.7% for two subsequent periods is about 50%. Thus, with probability 0.5 a region whose current deviation is large (larger than 0.7%), will move closer to the mean next year. As the first and fifth row of the upper panel of table 4 shows, there is a considerable probability that a region's inflation rate will not only switch to the adjacent state but to a state further away. Insofar, the reported figures are evidence of a strong within-distribution dynamics across U.S. metropolitan inflation rates. The extent of dynamics is even more pronounced for Japanese prefectures. From the number chosen to classify spaces, the lower degree of overall dispersion is clearly evident. Looking at the second panel of table 4, we can see that the probability of remaining in an "extreme" state is only about 30%. Thus, the likelihood of moving closer to the mean is 70%. Moreover, there is a considerable probability that existing mean deviations almost completely vanish. On the other hand, there is a significant dynamics away from the mean towards the tails of the distribution. As row three of the second panel shows, the probability of deviating by more than 0.3% from mean inflation next period when current deviation is below 0.1% is more than 30%. Overall, the results for Japan show that there is large within-distribution dynamics across prefectures. In terms of similarities in distribution patterns with the EMU case, the joint U.S./Canadian sample provides the best fit. As the third panel of table 4 shows the transition probability for the U.S./Canadian economic union has a comparable structure as that for the EMU (see table 8 of Weber and Beck (2003)). Particularly noteworthy is the fact, that there is less tendency for regions in the left or right tail of the distribution to exhibit large shifts towards the mean. Critics of EMU might take this evidence as support for their view that a single monetary policy is not adequate for the heterogeneous group of European countries. However, repeating arguments already mentioned above, this is not necessarily the case for a couple of reasons. First, U.S. and Canadian regions are likely relatively closely linked despite there is no single monetary policy. Besides NAFTA, i.e. economic linkages, there are close cultural and administrational links between the two countries. Additionally, Canadian monetary policies is not independent of the U.S. monetary policy but follows is more or less closely. Secondly, EMU is probably in a transition phase such that we expect within distribution dynamics to increase over time. Thirdly, given the evidence of dynamics towards the mean, it is the degree of overall dispersion that is important for policy makers. As we saw in section 4, overall dispersion in EMU is not extraordinary high such that our results do not raise serious doubts in the establishment and maintenance of a single monetary policy for EMU countries. As already outlined, the results for the transition probabilities rather are evidence of existing stronger rigidities in European markets.

In this and the previous sections, we have focused on analyzing the nature of the evolution of regional inflation dispersion for three different samples. Additionally, we have set the results in relation to the evidence for EMU. Now, we want to use the regional dimension of our data to perform a little exercise. This exercise will provide us with critical values for mean inflation rates below which a significant share of regions faces negative inflation. These values can be directly used by U.S. and Japanese policy makers as indicators for significant deflationary threats in their respective countries. However, they also have some benchmark character for EMU.

## 6 Mean Inflation Rates and Inflation Dispersion

In recent months, there has been a vivid discussion on how large the probability is that countries like the U.S. or particularly Germany might experience a deflation in the near future. In all contributions, the severe consequences of persistent negative inflation rates are emphasized. In a recent speech, Bernanke (2002), e.g., warns of the problem of "debt-deflation", i.e. an deflation induced, ever-increasing real value of debts, that can lead to fragility of a country's financial system. Ahearne et al. (2002) take the Japanese deflation experience in the 1990s to draw lessons for monetary policy in other countries potentially endangered by deflation. To prevent deflations various - and sometimes unorthodox - means are suggested. Bernanke (2002), e.g., suggests the following three measures: First, a central bank should preserve a buffer zone for inflation rates below which it should not push inflation. Secondly, central banks should forcefully ensure financial stability. And thirdly, when inflation rates are already low, central banks should act more aggressively than usual. In this section, we argue that regional dispersion in inflation rates provides an important aspect that a central bank facing the threat of a deflation has to bear into mind. As figures 1 and 2 have shown, regional inflation rates differ considerably. This can imply that some regions face negative inflation rates even if aggregate inflation rates are still well above 0. When dispersion is large, this can occur to a large extent. Then, debt holders in these regions may suffer from ever-creasing real values for their debts which might push local banks into trouble. Our strategy in examining this issue a little further is as follows: We will start by using the regional dimension of our data to approximate existing empirical inflation dispersions by a theoretical counterpart. This enables us to compute "critical values" for country-wide average inflation rates. These critical values indicate which proportion of all regions faces negative inflation rates given a certain national average inflation rate and given the assumed theoretical distribution. This approach allows us to assess how serious the problems risen above actually are.

Before proceeding with searching for an adequate theoretical distribution that fits the main characteristics of our data, we have to pay attention to an observation that we shortly mentioned above. The plots for the evolution of inflation rates (figures 1 to 3) together with the plots of the dynamics of overall dispersion (figures 11 to 13) indicate a relationship between the overall level of inflation and its degree of regional dispersion. This is particularly pronounced for the early period in the U.S. sample where the strong reduction in inflation is accompanied by a strong reduction in overall dispersion. If such a relationship turned out to be statistically significant but if we failed to take it into account in our computation of "critical values", we would overestimate the share of regions facing negative inflation for a given mean inflation rate.

Similar to our finding of a supposedly significant positive relationship between a country's average inflation rate and its regional inflation dispersion, a large strand of literature has empirically examined an analogous relationship between a country's inflation rate and its cross-sectional dispersion[12]. Theoretical models that try to explain this link can be mainly classified into two groups: menu-cost models (Sheshinski and Weiss (1977), Rotemberg (1983) and others) and signal extraction models (Lucas (1973), Barro (1976) and Hercowitz (1981)). Our results show that this relationship also has a regional dimension. It is easily conceivable that some of the mechanisms responsible for the link between the level of inflation and its variability across sectors generate a similar relationship between a country's average inflation rate and the cross-regional dispersion. Imagine, e.g., that price adjustments are costly. Then, local suppliers will adjust their prices not continuously but in steps, with the step size positively depending on the level of average inflation.

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[12] See e.g. Parks (1978), Fischer (1981) and Taylor (1981).

If price adjustment costs differ across regions or if there are region-specific shocks, staggered price setting across regions will occur and thus higher inflation will increase inflation dispersion across regions.

In search of an appropriate theoretical distribution necessary to compute "critical values" we particularly need to make sure that the left and right tails of the theoretical approximation capture their empirical counterparts sufficiently well. Our preferred candidate would be a normal distribution for several reasons. First, only first and second moments are required to fully describe the distribution. Secondly, computing critical values is very straightforward (see below). Fortunately, normal distributions seem to fit current inflation dispersions sufficiently well, at least for our purposes. Figure 21 plots empirical density estimates for inflation dispersions in 2001 for all three samples. Additionally, the plots contain the respective normal density functions used for approximation reasons. Although the fit is naturally not perfect, we think that it is good enough for the purpose of computing "critical" mean inflation rates values. Given our choice of normal distributions, we only need to compute first and second moments of empirical inflation dispersions to get theoretical approximations. Here, two things have to be observed. First, when computing the moments we want to pay attention to the different economic weights that the various regions have. We do this by weighting each region's inflation rate by its respective share in total GDP[13]. Secondly, we have to make sure that the supposed relationship between the average level of inflation and its cross-regional dispersion is taken into account. In other words, when determining the share of regions with negative inflation rates in dependence of the country-wide average inflation rate, we have take into account the fact that dispersion declines the closer average inflation approaches zero. We do this by estimating a functional relationship between the observed mean inflation and its cross-regional dispersion. More specifically, for each sample, we estimate an equation of the form

$$\sigma_t = \alpha + \beta * |\mu_t| + \epsilon_t, \tag{10}$$

where $\mu_t$ denotes the country-wide average inflation rate in period $t$ and $\sigma_t$ denotes the standard deviation of inflation dispersion in period $t$. The use of absolute values for mean inflation shows that we assume a symmetric relationship for positive and negative mean inflation rates. Certainly, this assumption may be problematic. However, as the number of deflationary periods is small even in the Japanese sample, this assumption can serve as a good working hypothesis, nevertheless. Estimation results

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[13] To compute weights, we are using national per capita GDP data from the OECD (2001 data). Weights are obtained by dividing the product of national per capita GDP data with a region's total population (obtained from the Bureau of Census for the U.S. and from http://www.population.de for Canada and Japan) through total GDP. Higher moments are computed using the same weights.

are presented in table 5. Both for the U.S. and the joint U.S./Canadian sample there exist a significant relationship between mean inflation and its regional dispersion. For Japan, however, this relationship is not significant. A possible explanation for these findings is that the suggested relationship between the mean and the dispersion of regional inflation rates is nonlinear. It is conceivable that the size of the relationship grows disproportionately with the mean or that there exist thresholds for the mean inflation rate below which the relationship is no longer existent. On the other hand, Japan-specific factors can also be a reason for the non-existence of such a relationship in this sample.

In the following, the computation of "critical values" is done from two slightly different perspectives. First, we will calculate "critical values" for mean inflation rates for which 1%, 2.5%, 5%, 10% and 25% of all regions face deflation. These computations are based on

$$\Phi\left(\frac{\pi - \mu_{crit}}{\sigma(\mu_{crit})}\right) = p_{crit}, \tag{11}$$

where $\Phi(.)$ denotes the cumulative density function of the normal distribution, $p_{crit}$ is the proportion of regions facing inflation rates below zero. $\mu_{crit}$ denotes the mean inflation rate associated with a share of $p_{crit}$ percent of all regions facing negative inflation rates. The expression $\sigma(\mu)$ reflects the relationship between mean inflation and inflation dispersion. To determine critical values for $\mu$, we set $\pi$ equal to zero and solve the equation for $\mu_{crit}$ using the estimation results from equation (10). This leads to:

$$\mu_{crit} = -\frac{\hat{\alpha} * \Phi^{-1}(p_{crit})}{1 + \hat{\beta} * \Phi^{-1}(p_{crit})}. \tag{12}$$

Results are presented in the upper panel of table 6. For comparison reasons, we also included previous findings for European regions (see Weber and Beck (2003)). The figures are very illustrative. Given the current degree of inflation dispersion, 5% of all regions in the U.S. will have inflation rates below 0 when the nation-wide inflation rate is about 1%. When national average inflation drops to 0.75%, the share of regions with inflation rates below zero raises to 10%. Thus, only for relatively low national-wide inflation rates, the share of regions with negative inflation rates becomes notable. Can we therefore totally ignore inflation dispersion when it comes to judging whether prevailing national inflation rates are posing deflationary threats? In our opinion, the answer is no. Given the findings of the Boskin report and the difficulties in adequately taking into account the effects of quality changes when assessing price changes, reported inflation rates probably overestimate the true inflation rate. Given rough estimates of this bias, a reported inflation rate of 1% corresponds to a *true* inflation rate of abount 0.25% to 0.5%. At these inflation rates, however, more than 20% of all regions have inflation rates below zero. In this sense, our findings provide an additional strong argument for Bernanke's buffer zone below which central banks should prevent inflation rates to fall. Analogous evidence for Japanese prefectures reflects our previous findings that regional dispersion in Japan is of minor size. Only when aggregate inflation is 0.5%, 10% of all regions have negative inflation rates. As we saw in the last sections, the case that is most comparable to EMU is that of the joint U.S./Canadian sample. As dispersion is larger for this sample and of comparable size to the EMU case, critical values are higher and of similar size as for EMU. However, despite a common monetary policy missing for the U.S. and Canada, critical values for EMU are generally higher. However, the 10% and 25% critical values are lower for the EMU. This "switch" reflects the much stronger relationship between mean and dispersion that we found for the European case. Thus, dispersion seems to decline stronger when mean inflation goes down. Then, of course, the share of regions falling into deflation increases less than proportionally with falling means.

The lower panel of table 6 takes a slightly different perspective. It reports shares of regions with inflation rates below zero in dependence of prevailing aggregate inflation rate. Thus, it gives a better intuition of how fast the share of "deflationary" regions increases with declining average inflation rates. With the exception of Japan, the general impression is that an aggregate inflation rate of 1% can be considered as establishing a "lower bound" below which inflation should not fall. When inflation falls below this point, the share of regions with negative inflation rates rises fast with each further tenth percentage point. Given the supposedly overestimation of actual inflation by reported numbers, a central bank should definitely not considering to push inflation below 1%. Interestingly, this finding is applicable for both the U.S. and the EMU as the comparison with our findings for Europe show.

# 7 Summary and Conclusions

In this paper, we used regional inflation rates for two well-established monetary unions and one economic union to analyze the nature of inflation dispersion in these samples. The importance of this question is evident. When inflation differences within a monetary union are large and persistent then the monetary authority will face contradicting demands: Whereas an expansionary policy might be adequate for the low-inflation region, a restrictionary policy is probably necessary for the high-inflation region. Thus, the question of inflation dispersion is of great relevance for monetary authorities. Research conducted in this paper has addressed this issue and has sought to answer the question of how large existing inflation dispersions are and how they evolve over time. Our regional data for the U.S., Canada and Japan show that inflation dispersion is considerable. Possible reasons include the existence of regionally segmented markets in conjunction with price-discriminating monopolists, different productivity trends across regions, short-run rigidities as well as asymmetric supply and demand shocks or a combination of all factors. Following the empirical growth literature, we used two basic approaches to examine how persistent inflation differentials are. Relying on the concept of $\beta$-convergence we found significant mean reversion in inflation rates. Estimation results indicate half lives of 6 months to slightly more than 1 year. The lowest half lives of around 6 months are found for Japan, for the U.S. we document half lives of around 9 months and for the U.S./Canadian sample half lives of slightly more than 1 year are found. Although these findings indicate some persistence in the factors causing inflation dispersion, half-lives reflect considerable dynamics towards the mean. Looking at the evolution of overall dispersion, we found strongly declining dispersion in the early 1980s and no or only minor declines afterwards. As there are considerable within distribution dynamics, the prevailing dispersion in inflation rates does not pose major problems to monetary authorities in the U.S or Japan. Such a conclusion is supported by our results in the section on distribution dynamics. Modeling the evolution of overall distribution dynamics as a time series process, we find for all samples indications of dynamics towards the cross-sectional mean. Again, the largest dynamics are found for Japan, followed by the U.S. and the joint U.S./Canadian sample. Estimated transition matrices indicate that there is a 50%-probability for a region with a large deviation from the cross-sectional mean to move closer towards the mean within one year. In an attempt to assess the significance of existing dispersion in regional inflation rates for the ongoing discussion on deflationary threats, we are providing "critical values" for aggregate inflation rates. In computing these values for U.S. data we are using a statistically significant relationship between mean inflation and its dispersion and approximate the empirical distribution by a normal density function. Our results indicate that aggregate inflation rates below 1% are associated with significant proportions of regions facing negative inflation rates. In connection with the likely overestimation of true inflation by reported number, we thus not only support Bernanke's buffer zone argument to prevent deflation but suggest a size of at least 1%.

One of our most striking results concerns the comparison of our findings in this paper with the evidence for EMU countries (see Weber and Beck (2003)). Both in terms of overall dispersion and within-distribution dynamics the best correspondence of EMU results with findings in this paper is found for the U.S./Canadian sample. This result is striking since the regions of the latter sample do not share a common monetary policy as the regions of EMU do. Depending on one's point of view, these findings can be interpreted differently. One certainly extreme position is that the evidence shows that EMU countries are not ready for a monetary union yet.

Obviously, there are strong asymmetric forces at work leading to a relatively large dispersion in inflation rates with relatively low tendencies towards reversion. While the fact of stronger rigidities in European inflation rates can certainly not be denied the conclusion that a single monetary policy for EMU is not adequate does not necessarily has to be drawn for several reasons. First, linkages between the U.S. and Canada are large with respect to many dimensions. There exists not only a free trade arrangement between these countries but there is a long history of close economic links. Additionally, the two countries share a common language and many other cultural and sociological characteristics. Additionally, there are close monetary linkages. Thus, the dispersion present in U.S./Canadian data cannot be expected to exhibit patterns of two economic entities that are subject to large asymmetric dynamics. Secondly, as the ECB has taken up its work only few years ago, the current state can certainly not be considered to be the steady state. It is very likely that there will be further steps towards convergence across member countries. Thirdly, even if there will be no more process towards further integration, an important question is whether the current extent of dispersion is unsustainable. Due to a lack of comparable events in the past, this question cannot be ultimately answered here. However, comparing absolute sizes of overall dispersion and existing convergence dynamics the figures for EMU do not seem to have unsustainable levels relative to U.S. figures, e.g. As we have shown in Weber and Beck (2003), there are significant tendencies towards the mean in EMU regional inflation rates. The difference mainly concerns the speed at which this occurs. As these differences very likely reflect large rigidities in markets across European countries, the ECB is definitely right when it asks politicians to do their homework and remove existing rigidities in European markets.

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# 8 Tables

Table 1: Countries and Regions/Cities included in our Study

| USA (24 metropolitan areas) |
|---|
| Anchorage, Atlanta, Boston, Chicago, Cincinnati, Cleveland, Dallas, Denver, Detroit, Honolulu, Houston, Kansas City, Los Angeles, Miami, Milwaukee, Minneapolis, New York, Philadelphia, Pittsburgh, Portland, St. Louis, San Diego, San Francisco, Seattle, |
| **Source**: Bureau of Labor Statistics<br>**Coverage**: 1980 - 2001 (CPI) |
| Canada (12 provinces) |
| Prince Edwards Islands, Alberta, New Brunswick, Nova Scotia, Quebec, Saskatchewan, New Foundland, Ontario, British Colombia, Yukon, Manitoba, Yellowknife |
| **Source**: CANSIM<br>**Coverage** 1980 - 2001 (CPI)<br>**Notes**: Annual inflation rates are computed as geometric means of monthly annual log changes of CPI index. For 2001, annual inflation rates were computed as geometric means from January till June inflation rates. For Yukon and Manitoba, data start in 1982. |
| Japan (47 prefectures) |
| Akita, Aomori, Chiba, Fukui, Fukuoka, Fukushima, Gifu, Hiroshima, Kagoshima, Kanazawa, Kobe, Kochi, Kofu, Kumamoto, Kyoto, Maebashi, Matsue, Matsuyama, Mito, Miyazaki, Morioka, Nagano, Nagasaki, Nagoya, Naha, Nara, Niigata, Oita, Okayama, Osaka, Otsu, Saga, Sapporo, Sendai, Shizuoka, Takamatsu, Tokushima, Ku-area of Tokyo, Tottori, Toyama, Tsu, Urawa, Utsunomiya, Wakayama, Yamagata, Yamaguchi, Yokohama |
| **Source**: Statistics Bureau and Statistics Center, Ministry of Public Management, Home Affairs, Post and Telecommunications<br>**Coverage** 1985 - 2001 (CPI)<br>**Notes**: Annual inflation rates are computed as geometric means of monthly annual log changes of CPI index. For 2001, annual inflation rates were computed as geometric means from January till April inflation rates. |

Table 2: Descriptive Statistics

| USA | | | | | |
|---|---|---|---|---|---|
| | 1981-2001 | 1981-1985 | 1986-1990 | 1991-1995 | 1996-2001 |
| mean | 3.62 | 5.36 | 3.69 | 3.07 | 2.59 |
| std.dvt. | 0.10 | 1.83 | 2.52 | 0.86 | 1.28 |

| Japan | | | | |
|---|---|---|---|---|
| | 1986-2001 | 1986-1990 | 1991-1995 | 1996-2001 |
| mean | 0.89 | 1.23 | 1.26 | 0.30 |
| std.dvt. | 0.04 | 0.60 | 0.29 | 0.36 |

| USA/Canada | | | | | |
|---|---|---|---|---|---|
| | 1981-2001 | 1981-1985 | 1986-1990 | 1991-1995 | 1996-2001 |
| mean | 3.64 | 5.84 | 3.76 | 2.83 | 2.34 |
| std.dvt. | 0.11 | 5.72 | 3.01 | 1.89 | 2.37 |

**Notes:**

1)The mean inflation rate (mean) is computed as the cross-sectional mean of geometric inflation means of the respectively included regions. The computation of the standard deviation is based on the cross-section of the regional geometric mean inflation rates.

2) Standard Deviations are multiplied by 10000.

Table 3: Panel Unit Root Tests (Levin and Lin (1993)) of Inflation Convergence

| Sample | $\rho$ | $\rho_{adj}$ | t-stat | p-value | half-live | half-live (adj.) |
|---|---|---|---|---|---|---|
| | | | USA | | | |
| 1981-2001 | 0.402 | 0.475 | -15.25 | 0.000 | 0.8 | 0.9 |
| 1981-1990 | 0.238 | 0.384 | -13.04 | 0.001 | 0.5 | 0.7 |
| 1991-2001 | 0.537 | 0.718 | -8.32 | 0.000 | 1.1 | 2.1 |
| | | | Japan | | | |
| 1986-2001 | 0.289 | 0.378 | -17.95 | 0.000 | 0.6 | 0.7 |
| 1991-2001 | 0.222 | 0.350 | -16.90 | 0.000 | 0.5 | 0.7 |
| | | | USA and Canada | | | |
| 1983-2001 | 0.451 | 0.527 | -17.07 | 0.000 | 0.9 | 1.1 |
| 1983-1990 | 0.450 | 0.697 | -12.62 | 0.011 | 0.9 | 1.9 |
| 1991-2001 | 0.525 | 0.703 | -10.64 | 0.000 | 1.1 | 2.0 |

**Notes:**
1) Results are based on the equation:

$$\Delta\tilde{\pi}_{i,t} = \alpha_i + \beta\tilde{\pi}_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j}\Delta\tilde{\pi}_{i,t-j} + \epsilon_{i,t},$$

where $\tilde{\pi}_{i,t}$ denotes the deviation of region's i inflation rate from the cross-sectional mean. A more detailed description of the estimation procedure is given in the appendix.

2) Bias adjustment is done using the formula given by Nickel (1981), see section B of the appendix.

Table 4: Transition Probabilities - Annual Transitions - for Deviations from Cross-Regional Mean

Transition Probabilities for USA

| Dev. in $t$ | Dev. in $t+1$ | | | | |
|---|---|---|---|---|---|
| | $< -0.70$ | $-0.20$ | 0.20 | 0.70 | $> 0.70$ |
| $< -0.70$ | 0.43 | 0.22 | 0.13 | 0.14 | 0.08 |
| $-0.20$ | 0.25 | 0.28 | 0.26 | 0.15 | 0.07 |
| 0.20 | 0.15 | 0.33 | 0.21 | 0.21 | 0.09 |
| 0.70 | 0.05 | 0.19 | 0.31 | 0.35 | 0.10 |
| $> 0.70$ | 0.02 | 0.11 | 0.09 | 0.21 | 0.56 |

Transition Probabilities for Japan

| Dev. in $t$ | Dev. in $t+1$ | | | | |
|---|---|---|---|---|---|
| | $< -0.30$ | $-0.10$ | 0.10 | 0.30 | $> 0.30$ |
| $< -0.30$ | 0.29 | 0.27 | 0.25 | 0.11 | 0.07 |
| $-0.10$ | 0.19 | 0.28 | 0.20 | 0.19 | 0.14 |
| 0.10 | 0.19 | 0.22 | 0.24 | 0.22 | 0.13 |
| 0.30 | 0.14 | 0.18 | 0.21 | 0.28 | 0.19 |
| $> 0.30$ | 0.09 | 0.21 | 0.18 | 0.19 | 0.33 |

Transition Probabilities for USA and Canada

| Dev. in $t$ | Dev. in $t+1$ | | | | |
|---|---|---|---|---|---|
| | $< -0.70$ | $-0.20$ | 0.20 | 0.70 | $> 0.70$ |
| $< -0.70$ | 0.48 | 0.24 | 0.11 | 0.09 | 0.07 |
| $-0.20$ | 0.25 | 0.22 | 0.28 | 0.16 | 0.09 |
| 0.20 | 0.19 | 0.26 | 0.18 | 0.19 | 0.18 |
| 0.70 | 0.08 | 0.12 | 0.29 | 0.21 | 0.30 |
| $> 0.70$ | 0.07 | 0.08 | 0.16 | 0.20 | 0.49 |

**Notes:**

1) Table entries report conditional probabilities for a region's inflation rate to transit from the inflation range indicated in the first column to the inflation range indicated in columns two to six. Ranges are defined in terms of deviations from the cross-regional mean. $-0.20$, e.g., includes all observations where the cross-regional-mean-inflation deviation in period $t$ is in the range $[-0.70, 0.20[$. Ranges were chosen such that each range has approximately the same number of observations.

Table 5: Examining the Relationship between Mean Inflation and Regional Inflation Dispersion

Estimated Equation: $\sigma_t = \alpha + \beta\mu_t + \epsilon_t$

| $\alpha$ | $\beta$ | $R^2_{adj}$ | $s.e.$ |
|---|---|---|---|
| USA | | | |
| 0.005<br>(0.001) | 0.066<br>(0.031) | 0.15 | 0.002 |
| Japan | | | |
| 0.004<br>(0.0003) | -0.020<br>(0.018) | 0.01 | 0.001 |
| USA and Canada | | | |
| 0.005<br>(0.002) | 0.094<br>(0.038) | 0.20 | 0.003 |

**Notes:**

1) $\sigma_t$ denotes the standard deviation of the empirical inflation distribution in period $t$, $\mu_t$ denotes its mean.

2) Numbers in brackets denoted standard deviations.

Table 6: Relationship between Average Inflation Rate and Proportion of Regions Facing Negative Inflation

”Critical” Average Inflation Rates

| Prop. of ”Defl.” Regions | USA | Japan | USA/Canada | EMU |
|---|---|---|---|---|
| 1% | 1.47 | 0.89 | 1.62 | 1.87 |
| 2.5% | 1.20 | 0.76 | 1.31 | 1.42 |
| 5% | 0.99 | 0.64 | 1.06 | 1.09 |
| 10% | 0.75 | 0.50 | 0.79 | 0.78 |
| 25% | 0.38 | 0.27 | 0.39 | 0.36 |

Mean Inflation Rate and Percentage of Regions with Deflation

| Mean Infl. Rate | USA | Japan | USA/Canada | EMU |
|---|---|---|---|---|
| 2.00 | 0.13 | 0.00 | 0.31 | 0.79 |
| 1.90 | 0.20 | 0.00 | 0.43 | 0.95 |
| 1.80 | 0.29 | 0.00 | 0.58 | 1.15 |
| 1.70 | 0.42 | 0.00 | 0.78 | 1.40 |
| 1.60 | 0.62 | 0.00 | 1.06 | 1.71 |
| 1.50 | 0.89 | 0.00 | 1.43 | 2.10 |
| 1.40 | 1.27 | 0.01 | 1.91 | 2.58 |
| 1.30 | 1.80 | 0.03 | 2.55 | 3.18 |
| 1.20 | 2.52 | 0.08 | 3.38 | 3.94 |
| 1.10 | 3.49 | 0.19 | 4.46 | 4.89 |
| 1.00 | 4.79 | 0.45 | 5.85 | 6.09 |
| 0.90 | 6.48 | 0.96 | 7.61 | 7.59 |
| 0.80 | 8.65 | 1.92 | 9.82 | 9.47 |
| 0.70 | 11.40 | 3.57 | 12.55 | 11.83 |
| 0.60 | 14.80 | 6.21 | 15.89 | 14.76 |
| 0.50 | 18.92 | 10.12 | 19.88 | 18.38 |
| 0.40 | 23.80 | 15.51 | 24.59 | 22.81 |
| 0.30 | 29.43 | 22.45 | 30.01 | 28.14 |
| 0.20 | 35.76 | 30.78 | 36.12 | 34.47 |
| 0.10 | 42.68 | 40.13 | 42.83 | 41.78 |
| 0.00 | 50.00 | 50.00 | 50.00 | 50.00 |

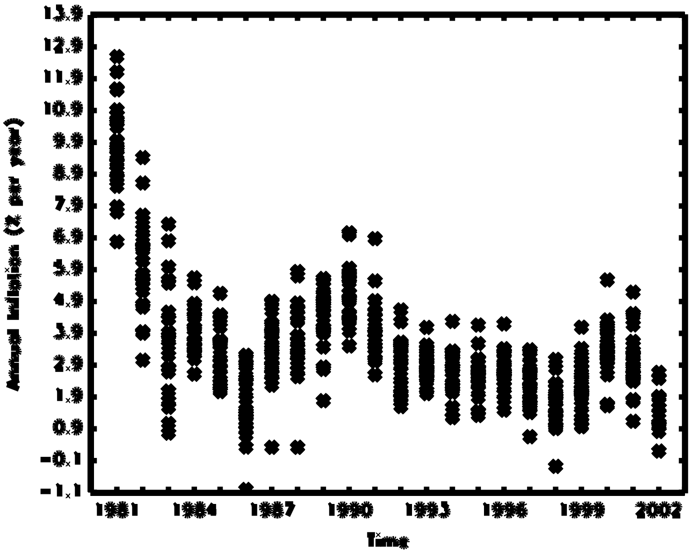
**Notes:**
1) Mean inflation rates are computed by weighting each regional inflation rate with the respective region’s share in total GDP, i.e.

$$\hat{\pi}_t = \sum_{i=1}^{N} \gamma_i \pi_{i,t},$$

with $\gamma_i = \frac{GDP_i}{\sum_{i=1}^{N}}$ denoting region’s i share in total gross domestic product (GDP).

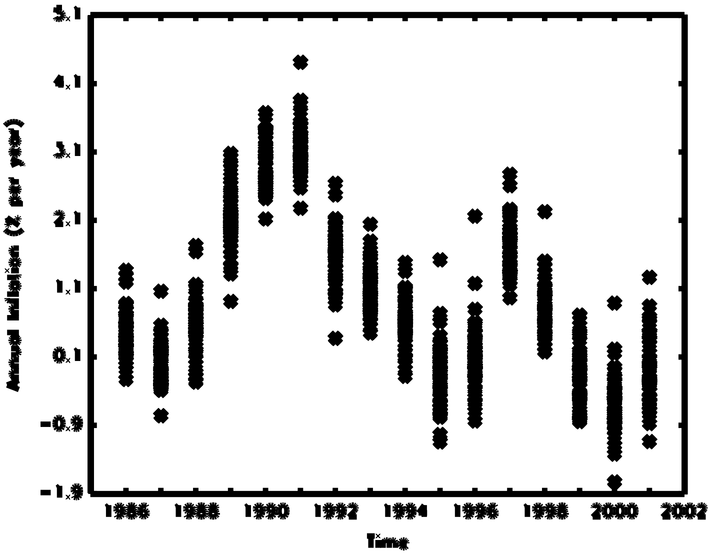
# 9 Figures

Figure 1: Inflation Dispersion Across U.S. Metropolitan Areas



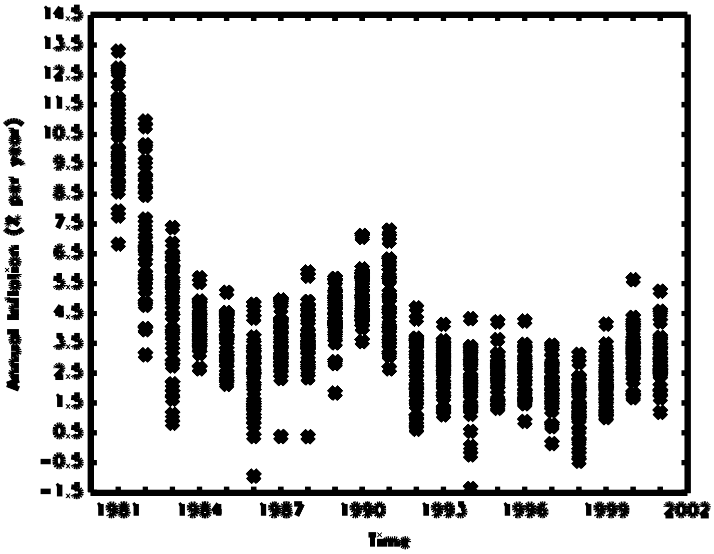
Note: *Figure 1 plots cross-sectional inflation rates for U.S. metropolitan areas. Inflation are computed as annual percentage changes*

Figure 2: Inflation Dispersion Across Japanese Prefectures



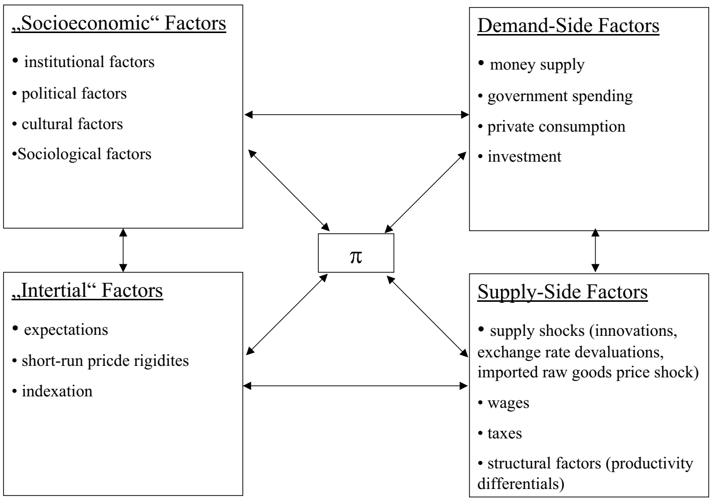
Note: *Figure 2 plots cross-sectional inflation rates for Japanese prefectures. Inflation are computed as annual percentage changes*

Figure 3: Inflation Dispersion Across U.S. Metropolitan Areas and Canadian Provinces



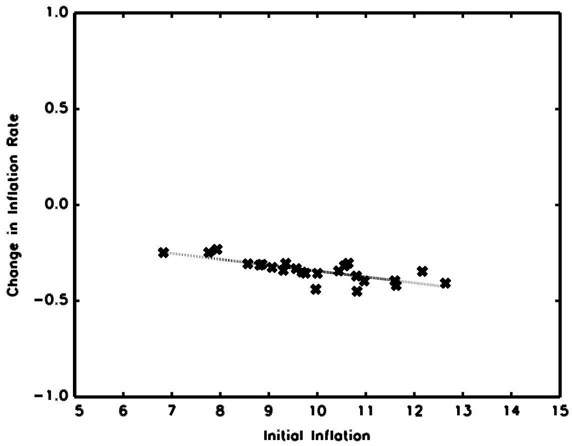
Note: *Figure 3 plots cross-sectional inflation rates for U.S. metropolitan areas and Canadian provinces. Inflation are computed as annual percentage changes*

Figure 4: Inflation Determinants: Overview



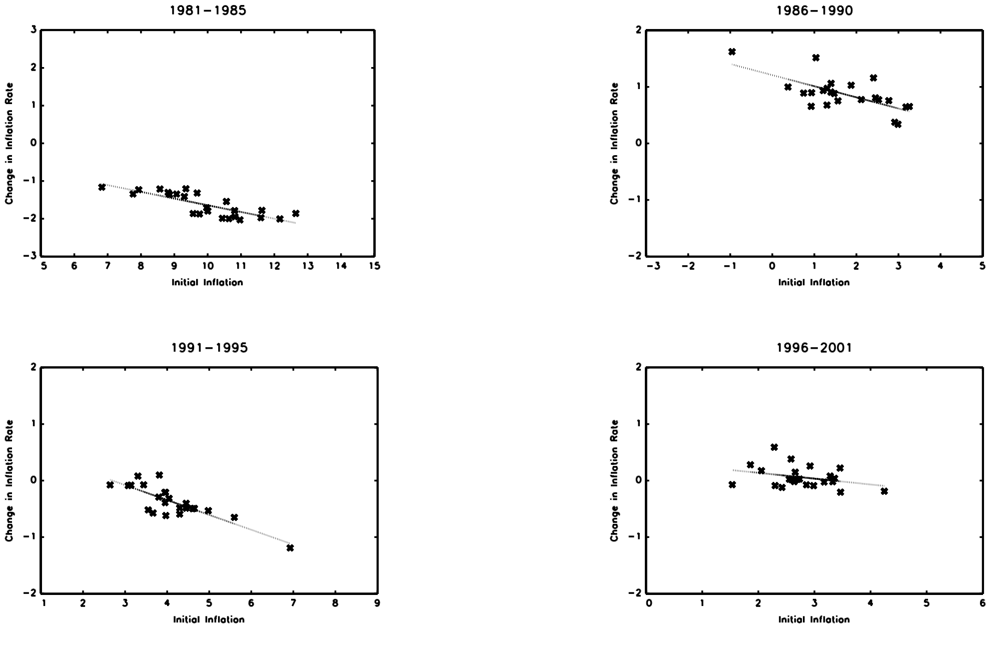
Note: *Figure 4 gives an overview of possible inflation determinants discussed in the literature on inflation. The individual factors are grouped into four categories. As the arrows indicate, individual factors are assumed to be interdependent.*

Figure 5: Change in Inflation vs. Initial Inflation: USA, Total Period



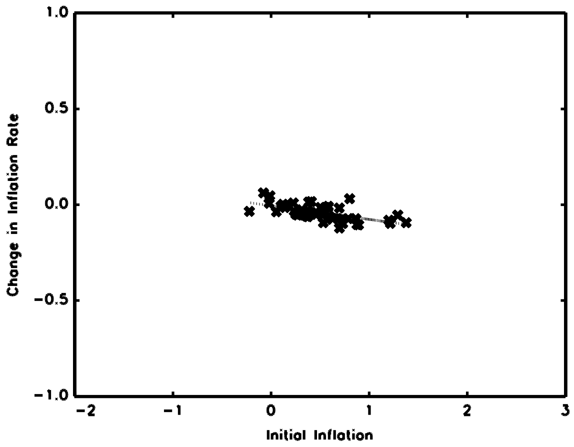
Note: *Figure 5 plots annual average changes in inflation rates ("All Items") for U.S. metropolitan areas versus annual average initial inflation rates. Inflation rates are computed as annual percentage changes. The dotted line plots fitted values from a OLS regression.*

Figure 6: Change in Inflation vs. Initial Inflation: USA, Subperiods



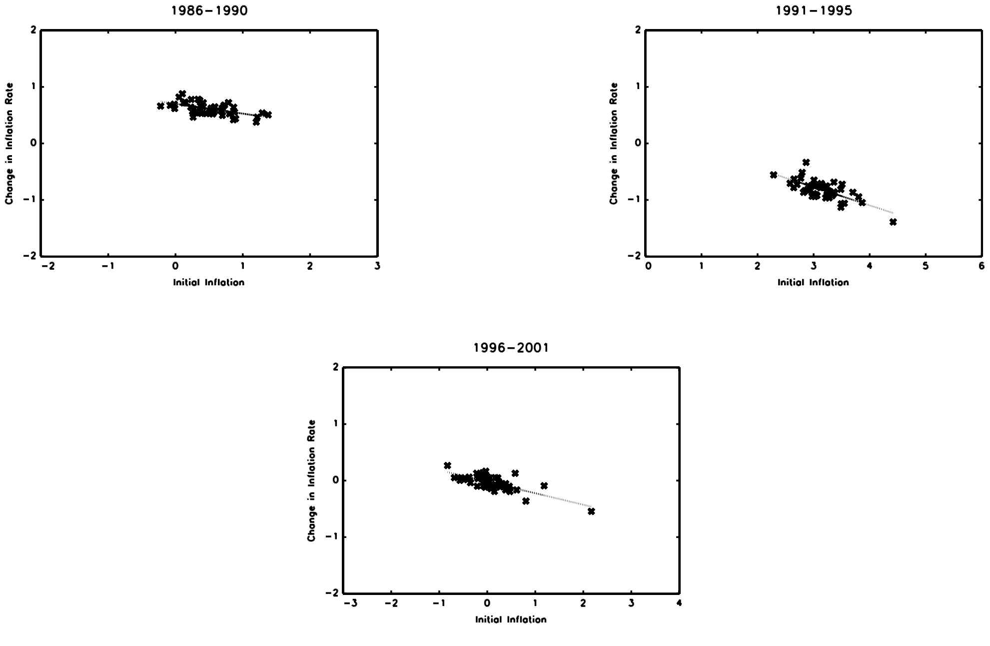
Note: *Figure 6 plots annual average changes in inflation rates ("All Items") for U.S. metropolitan areas versus annual average initial inflation rates for four subperiods. Inflation rates are computed as annual percentage changes. The dotted line plots fitted values from a OLS regression.*

Figure 7: Change in Inflation vs. Initial Inflation: Japan, Total Period



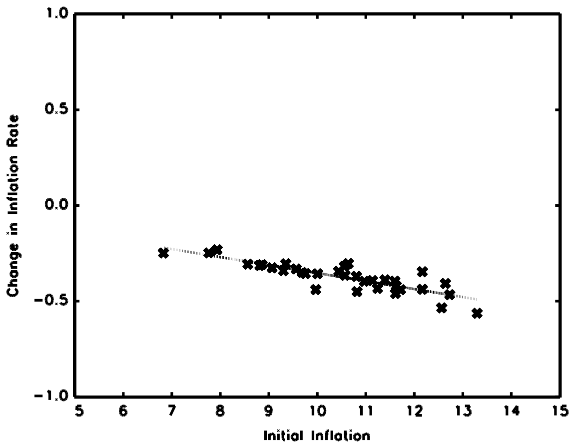
Note: *Figure 7 plots annual average changes in inflation rates ("All Items") for Japanese prefectures versus annual average initial inflation rates. Inflation rates are computed as annual percentage changes. The dotted line plots fitted values from a OLS regression.*

Figure 8: Change in Inflation vs. Initial Inflation: Japan, Subperiods



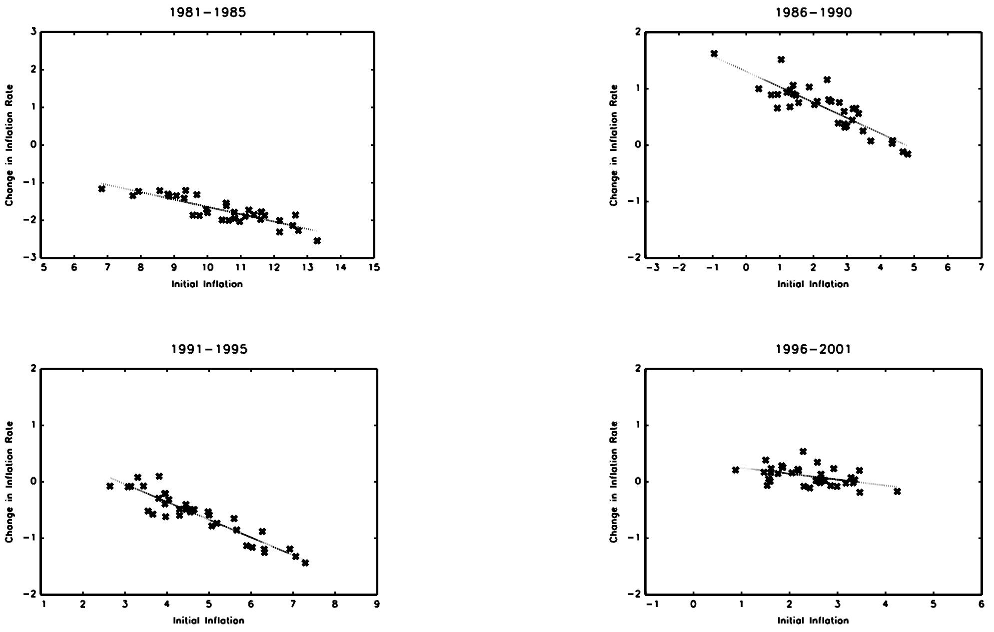
Note: *Figure 8 plots annual average changes in inflation rates ("All Items") for Japanese prefectures versus annual average initial inflation rates for three subperiods. Inflation rates are computed as annual percentage changes. The dotted line plots fitted values from a OLS regression.*

Figure 9: Change in Inflation vs. Initial Inflation: USA AND CANADA, Total Period



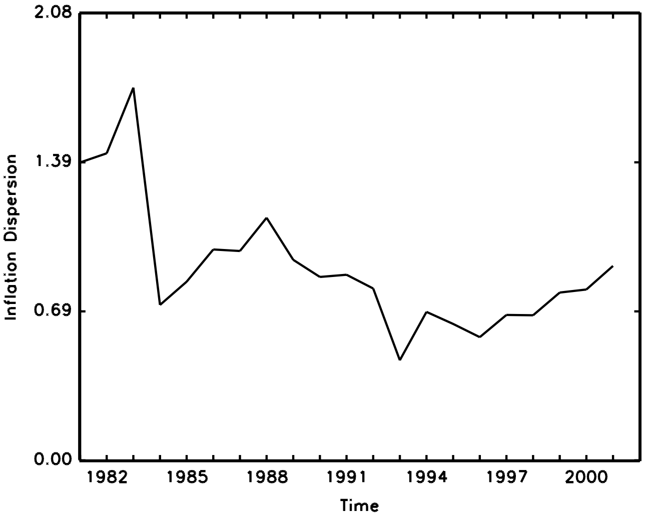
Note: *Figure 9 plots annual average changes in inflation rates ("All Items") for U.S. metropolitan areas and Canadian provinces versus annual average initial inflation rates. Inflation rates are computed as annual percentage changes. The dotted line plots fitted values from a OLS regression.*

Figure 10: Change in Inflation vs. Initial Inflation: USA AND CANADA, Subperiods



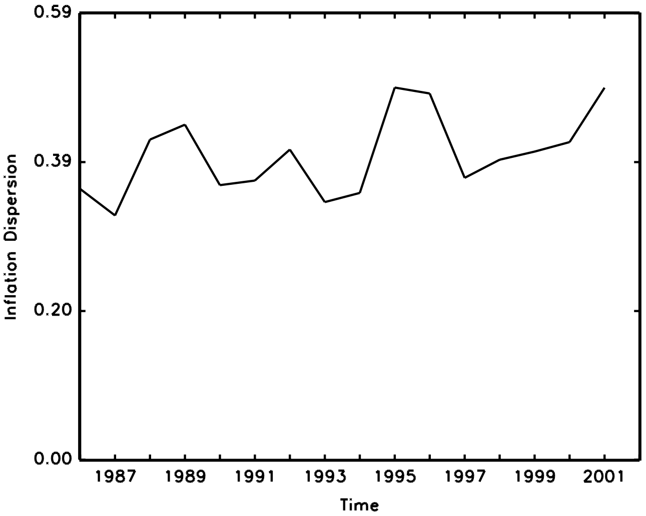
Note: *Figure 10 plots annual average changes in inflation rates ("All Items") for U.S. metropolitan areas and Canadian provinces versus annual average initial inflation rates for four subperiods. Inflation rates are computed as annual percentage changes. The dotted line plots fitted values from a OLS regression.*

Figure 11: Cross-Regional Inflation Dispersion: USA, Total Period



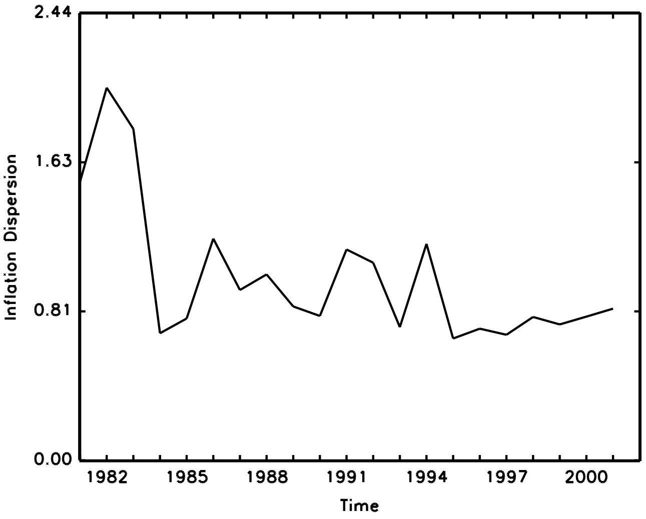
Note: *Figure 11 plots the standard deviation of regional inflation dispersion across U.S. metropolitan areas. Inflation rates are computed as annual percentage changes in the respective index. All figures are multiplied by 100.*

Figure 12: Cross-Regional Inflation Dispersion: Japan, Total Period



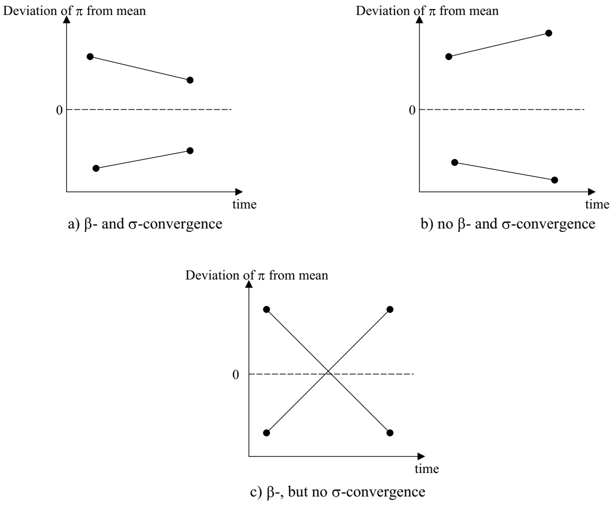
Note: *Figure 12 plots the standard deviation of regional inflation dispersion across Japanese prefectures. Inflation rates are computed as annual percentage changes in the respective index. All figures are multiplied by 100.*

Figure 13: Cross-Regional Inflation Dispersion: USA and Canada, Total Period



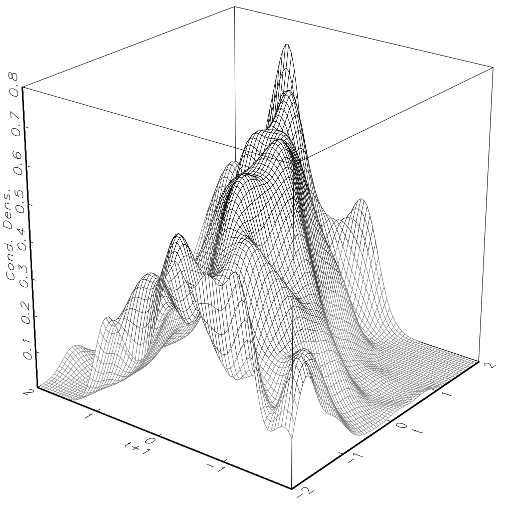
Note: *Figure 13 plots the standard deviation of regional inflation dispersion across U.S. metropolitan areas and Canadian provinces. Inflation rates are computed as annual percentage changes in the respective index. All figures are multiplied by 100.*

Figure 14: The Relationship between $\beta$- and $\sigma$-Convergence



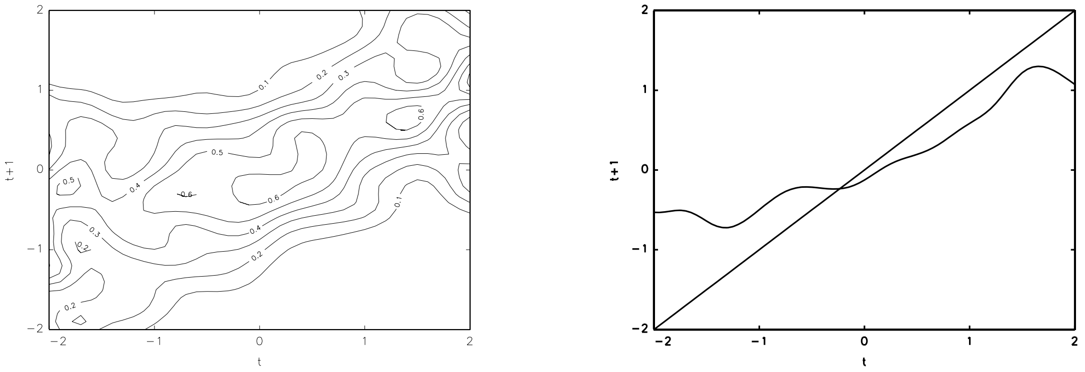
Note: *Figure 14 illustrates the relation between $\beta$- and $\sigma$-convergence. The individual panels thereby reflect three different possibilities of the evolution of the inflation of two different regions is plotted.*

Figure 15: Stochastic Kernel: USA, Annual Transitions



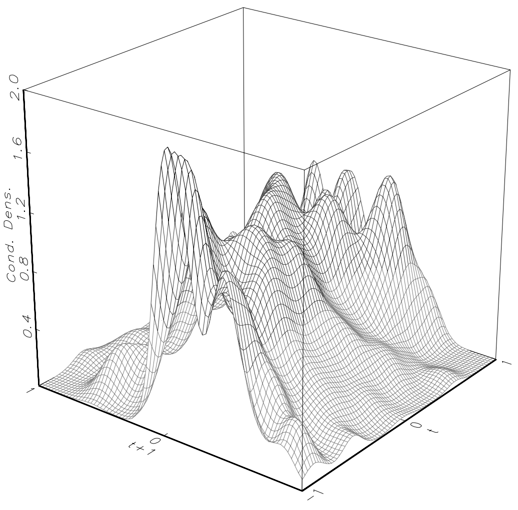
Note: *Figure 15 represents the surface plot of the stochastic kernel for annual-inflation-rate transitions (1983 to 2001) for U.S. metropolitan ares. On the x-axis (denoted by t), period's t inflation deviations from the cross-regional mean and on the y-axis (denoted by* $t+1$*), period's* $t+1$ *inflation deviations from the cross-regional mean are plotted. On the z-axis, the transition density function* $p(x,y)$ *associated with the stochastic kernel* $P(x,A)$ *is plotted*

Figure 16: Stochastic Kernel: USA, Annual Transitions, Contour Plot and Conditional Expected Mean Inflation Deviations



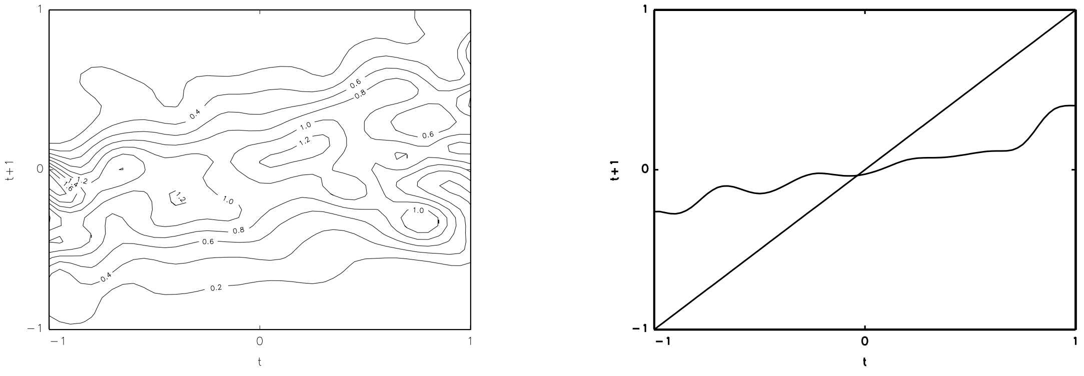
Note: *The left panel of figure 16 represents the contour plot of the transition density function* $p(x,y)$ *associated with the stochastic kernel* $P(x,A)$ *that we computed for the U.S. metropolitan areas (see figure 15). The right panel of figure 16 shows how corresponding conditional expected period's* $t+1$ *mean-inflation deviations behave relative to period's t mean-inflation deviations.*

Figure 17: Stochastic Kernel: Japan, Annual Transitions



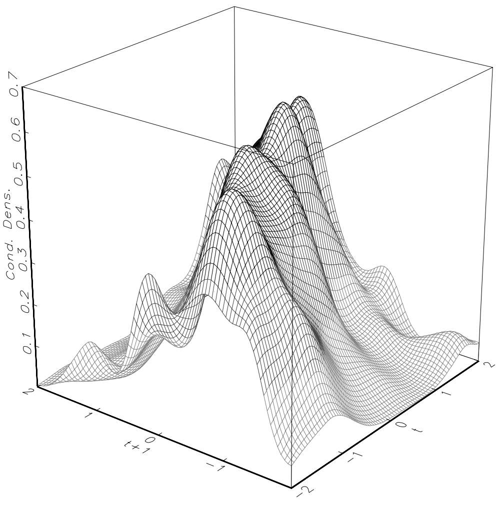
Note: *Figure 17 represents the surface plot of the stochastic kernel for annual-inflation-rate transitions (1986 to 2001) for Japanese prefectures. On the x-axis (denoted by t), period's t inflation deviations from the cross-regional mean and on the y-axis (denoted by $t + 1$), period's $t + 1$ inflation deviations from the cross-regional mean are plotted. On the z-axis, the transition density function $p(x, y)$ associated with the stochastic kernel $P(x, A)$ is plotted*

Figure 18: Stochastic Kernel: Japan, Annual Transitions, Contour Plot and Conditional Expected Mean Inflation Deviations



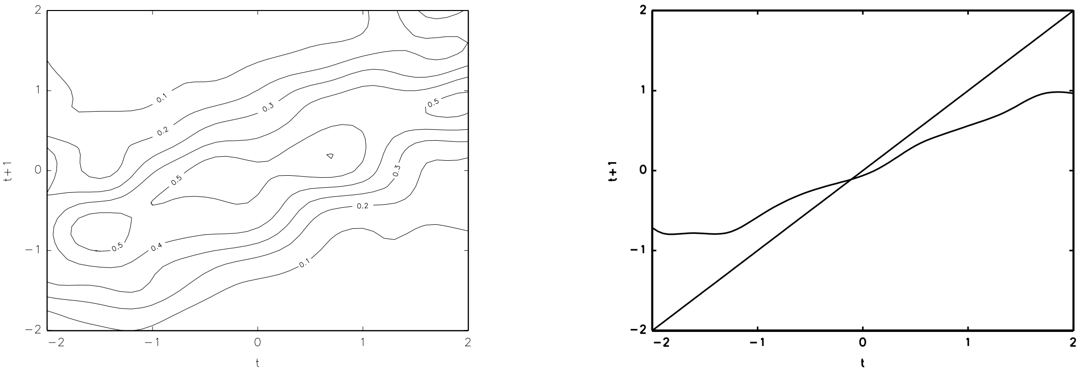
Note: *The left panel of figure 18 represents the contour plot of the transition density function $p(x, y)$ associated with the stochastic kernel $P(x, A)$ that we computed for Japanese prefectures (see figure 17). The right panel of figure 18 shows how corresponding conditional expected period's $t + 1$ mean-inflation deviations behave relative to period's t mean-inflation deviations.*

Figure 19: Stochastic Kernel: USA and Canada, Annual Transitions



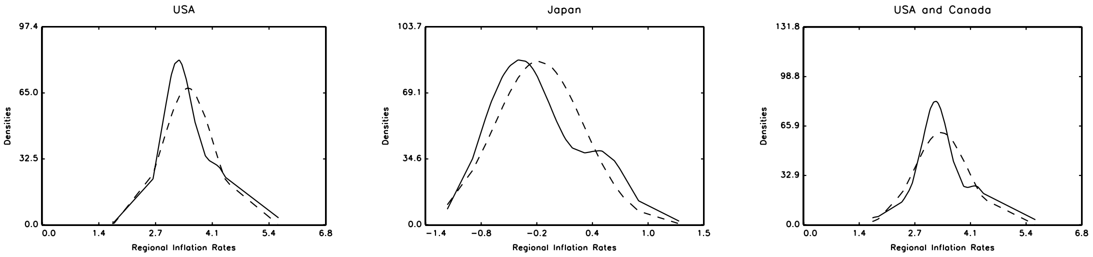
Note: *Figure 19 represents the surface plot of the stochastic kernel for annual-inflation-rate transitions (1986 to 2001) for U.S. metropolitan areas and Canadian provinces. On the x-axis (denoted by t), period's t inflation deviations from the cross-regional mean and on the y-axis (denoted by* $t+1$*), period's* $t+1$ *inflation deviations from the cross-regional mean are plotted. On the z-axis, the transition density function* $p(x, y)$ *associated with the stochastic kernel* $P(x, A)$ *is plotted*

Figure 20: Stochastic Kernel: USA and Canada, Annual Transitions, Contour Plot and Conditional Expected Mean Inflation Deviations



Note: *The left panel of figure 20 represents the contour plot of the transition density function* $p(x, y)$ *associated with the stochastic kernel* $P(x, A)$ *that we computed for U.S. metropolitan areas and Canadian provinces (see figure 19). The right panel of figure 20 shows how corresponding conditional expected period's* $t + 1$ *mean-inflation deviations behave relative to period's t mean-inflation deviations.*

Figure 21: Empirical Density Functions of Regional Inflation Dispersion and Theoretical Approximations



Note: *Figure 21 plots kernel density estimates of empirical regional inflation distributions versus the density from a normal distribution used as an approximation. The empirical distribution is that prevailing in 2000. The left panel plots data for the U.S. sample, the medium panel plots data for Japanese prefectures and the right panel plots data for the U.S. and Canadian sample.*

# 10 Appendix

## A The Levin-Lin Panel-Unit-Root Test

Our panel-unit-root analysis in section 3 is based on Levin and Lin (1993). To compute the panel-unit root results, we proceed as follows: Let $\{\pi_{i,t}\}$ be a balanced panel of inflation rates consisting of $N$ individual regions with $T$ observations, respectively. We assume that the individual observations are generated by

$$\Delta \pi_{i,t} = \rho_i \pi_{i,t-1} + u_{i,t} \tag{13}$$

where $-2 < \rho_i \leq 0$, and $u_{i,t}$ has the following error-components representation

$$u_{i,t} = \alpha_i + \theta_t + \epsilon_{i,t}. \tag{14}$$

In this specification, $\alpha_i$ represents an individual-specific effect, $\theta_t$ represents a common-time effect and $\epsilon_{i,t}$ is a (possibly serially correlated) stationary idiosyncratic shock. The Levin-Lin test procedure imposes (both for the null hypothesis of nonstationarity and for the alternative hypothesis of stationarity) the homogeneity restriction that all $\beta$s are equal across individual regions. Thus, the null hypothesis can be formulated as:

$$H_0 : \rho_1 = \rho_2 = \cdots = \rho_N = \rho = 0,$$

and the alternative hypothesis (that all series are stationary) is given by:

$$H_1 : \rho_1 = \rho_2 = \cdots = \rho_N = \rho < 0,$$

To test the null hypothesis of nonstationarity we proceed as follows:

1. First, we control for the common-time effect by subtracting the cross-sectional means:

$$\tilde{\pi}_{i,t} = \pi_{i,t} - \frac{1}{N} \sum_{j=1}^{N} \pi_{j,t} \tag{15}$$

Having transformed the dependent variable in this way our test equation becomes:

$$\Delta \tilde{\pi}_{i,t} = \rho \tilde{\pi}_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j} \Delta \tilde{\pi}_{i,t-j} + \epsilon_{i,t}. \tag{16}$$

To control for potential serial correlations in the idiosyncratic shocks $\epsilon_{i,t}$ we include lagged differences of $\tilde{\pi}_{i,t}$. Whereas we equalize the $\rho$s across individuals we allow for different degrees of serial correlation $k_i$ (with $i = 1, \ldots, N$) across them. The number of lagged differences for each region is determined by Hall (1994) general-to-specific method, recommended by Campbell and Perron (1991).

2. The next step in our testing procedure is to run the following two auxiliary regressions
$$\Delta\tilde{\pi}_{i,t} = \sum_{j=1}^{k_i} \phi_{1i,j}\Delta\tilde{\pi}_{i,t-j} + e_{i,t}. \tag{17}$$
$$\tilde{\pi}_{i,t-1} = \sum_{j=1}^{k_i} \phi_{2i,j}\Delta\tilde{\pi}_{i,t-j} + \nu_{i,t-1}. \tag{18}$$
and to retrieve the residuals $\hat{e}_{i,t}$ and $\hat{\nu}_{i,t-1}$ from these regressions.

3. These residuals are used to run the regression
$$\hat{e}_{i,t} = \rho_i\hat{\nu}_{i,t-1} + \eta_{i,t}. \tag{19}$$
The residuals of (19) are used to compute an estimate of the variance of $\eta_{i,t}$:
$$\hat{\sigma}^2_{ei} = \frac{1}{T-k_i-1}\sum_{t=k_i+2}^{T} \hat{\eta}^2_{i,t} \tag{20}$$

4. Normalizing the OLS residuals $\hat{e}_{i,t}$ and $\hat{\nu}_{i,t-1}$ by dividing them through $\hat{\sigma}_{ei}$ yields:
$$\tilde{e}_{i,t} = \frac{\hat{e}_{i,t}}{\hat{\sigma}_{ei}} \tag{21}$$
$$\tilde{\nu}_{i,t-1} = \frac{\hat{\nu}_{i,t-1}}{\hat{\sigma}_{ei}} \tag{22}$$

5. The normalized residuals are used to run the following pooled cross-section time-series regression:
$$\tilde{e}_{i,t} = \rho\tilde{\nu}_{i,t-1} + \tilde{\epsilon}_{i,t}. \tag{23}$$
Under the null hypothesis,- $\tilde{e}_{i,t}$ is independent of $\tilde{\nu}_{i,t-1}$, i.e. we can test the null hypothesis by testing whether $\delta = 0$. Unfortunately, the studentized coefficient
$$\tau = \frac{\hat{\rho}}{\hat{\sigma}_{\tilde{\epsilon}}\sum_{i=1}^{N}\sum_{t=2+k_i}^{T}\tilde{\nu}^2_{i,t-1}}$$
with
$$\hat{\sigma}_{\tilde{\epsilon}} = \frac{1}{NT}\sum_{i=1}^{N}\sum_{t=2+k_i}^{T}\tilde{\epsilon}_{i,t}$$

is not asymptotically normally distributed. Levin and Lin (1993) compute an adjusted test statistic based on $\tau$ that it is asymptotically normally distributed. However, we do not make use of their adjustment procedure but use bootstrap methods to compute critical values for the null hypothesis. This procedure is described in Section C.

## B Bias Adjustment of the Panel Root Estimate

To compute the half-live to convergence of deviations from the law of one price we use the formula $t_{half} = \frac{-\ln(2)}{ln(\rho)}$ where $\rho$ is described by the following relationship: $\rho = \rho - 1$. Unfortunately, the estimated value of $\rho$, denoted $\hat{\rho} = \hat{\rho} + 1$ is biased downward and has therefore to be adjusted before the half-live to convergence can be computed. We use the formula by Nickel (1981) who showed that

$$\underset{\mathrm{N}\longrightarrow\infty}{\mathrm{plim}}\,(\hat{\rho} - \rho) \longrightarrow \frac{A_T B_T}{C_T}$$

with $A_T = \frac{-(1+\rho)}{T-1}$, $B_T = 1 - \frac{1}{T}\frac{(1-\rho^T)}{(1-\rho)}$ and $C_T = 1 - \frac{2\rho(1-B_T)}{[(1-\rho)(T-1)]}$.

## C Parametric Bootstrap

To derive critical values for our null hypothesis that $\rho$is equal to zero, we employ a parametric bootstrap procedure that proceed as follows:

1. We start by estimating the data generating process (DGP) under the null hypothesis, given by

$$\Delta\pi_{i,t} = \sum_{j=1}^{k_i} \phi_{i,j}\Delta\pi_{i,t-j} + \epsilon_{i,t} \tag{24}$$

.

2. Then, we draw a sample of T + R innovation vectors from a standardized normal distribution that we use to recursively build up a panel of pseudo-observations $\pi_{i,t_{i=1,2,\ldots,N;t=1,2,\ldots,T}}$ based on the estimated GDP.

3. Then - after dropping the first R observations - we perform the Levin and Lin (1993) test (as described in section A) on these observations (without subtracting the cross-sectional mean). The resulting studentized coefficient $\tau$ is saved.

4. Steps 2 and 3 are repeated many times (5000). The collection of the $\tau$ statistics form the bootstrap distribution of these statistics under the null hypothesis.