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Guenter W. Beck and Axel A. Weber

Center for Financial Studies
an der J. W. Goethe-Universität ▪ Taunusanlage 6 ▪ D-60329 Frankfurt am Main
Telefon: (+49)069/242941-0
Fax: (+49)069/242941-77 ▪ http://www.ifk-cfs.de ▪ E-Mail: info@ifk-cfs.de

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Center for Financial Studies
an der J. W. Goethe-Universität ■ Taunusanlage 6 ■ D-60329 Frankfurt am Main
Telefon: (+49)069/242941-0
Fax: (+49)069/242941-77 ■ http://www.ifk-cfs.de ■ E-Mail: info@ifk-cfs.de

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Inflation Rate Dispersion and Convergence in Monetary and Economic Unions: Lessons for the ECB*

Guenter W. Beck[1] and Axel A. Weber[2]

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Abstract:
Using a set of regional inflation rates we examine the dynamics of inflation dispersion within the U.S.A., Japan and across U.S. and Canadian regions. We find that inflation rate dispersion is significant throughout the sample period in all three samples. Based on methods applied in the empirical growth literature, we provide evidence in favor of significant mean reversion (β-convergence) in inflation rates in all considered samples. The evidence on σ-convergence is mixed, however. Observed declines in dispersion are usually associated with decreasing overall inflation levels which indicates a positive relationship between mean inflation and overall inflation rate dispersion. Our findings for the within-distribution dynamics of regional inflation rates show that dynamics are largest for Japanese prefectures, followed by U.S. metropolitan areas. For the combined U.S.-Canadian sample, we find a pattern of within-distribution dynamics that is comparable to that found for regions within the European Monetary Union (EMU). In line with findings in the so-called 'border literature' these results suggest that frictions across European markets are at least as large as they are, e.g., across North American markets.

JEL Classification: E31, E52, E58

Keywords: Inflation Convergence, Deflation, ECB Monetary Policy, EMU, Regional Diversity

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[1] Corresponding Author, University of Frankfurt and CFS, phone: ++49-69-79828320, email: gbeck@wiwi.uni-frankfurt.de

[2] Deutsche Bundesbank

1 Introduction

One of the major challenges that the European Central Bank (ECB) faces is the extent and dynamics of inflation rate 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). One issue that has been paid particular attention to in this context are the implications of Balassa-Samuelson effects for inflation rate 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 those in 'richer' member countries such as France or Germany. The resulting inflation rate dispersion constitutes a large problem for the ECB: When it strictly sticks to its target (an EMU-wide average inflation rate of less than (but close to) 2%), several countries might face the danger of persistently low or even negative inflation rates. However, when it tolerates an EMU-wide average inflation rate 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 rate dispersion and convergence has played only a minor role in the literature before the establishment of the EMU. One reason for this neglect might be that in well-established monetary unions such as the U.S.A, Japan or Germany, regional inflation rate dispersion was not present or was of negligible size. That this is not the case is clearly shown by Cecchetti et al. (2002) 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 rate dispersion might be that all well-established monetary unions were taken as a datum and were considered to be more or less optimum currency areas. Thus, research in monetary policy focussed less on inflation rate dispersion within monetary unions but on other aspects such as strategic interactions between policy-makers and the public, the correct handling of uncertainties, etc. The view of existing monetary unions as optimum currency areas is challenged by Rockoff (2000) among others, 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 been positive." According 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, an assessment of current or potential future developments in the euro area is difficult. One way to overcome this problem is to refer to evidence from well-established monetary unions. In our opinion, such an approach can be helpful for at least two reasons. First, a comparison with prevailing inflation rate dispersion within the U.S., e.g., can be helpful to get a better assessment of the extent of existing inflation rate dispersion in the euro area. We particularly want to

get some intuition about the question whether EMU-wide inflation rate dispersion appears to be of sustainable size when compared to evidence from other countries. Secondly, the study of long existing (but not necessarily optimum) monetary unions can give us a hint where EMU inflation rate dispersion might evolve over time. Based on this reasoning, this paper considers dispersion evidence of two monetary unions (U.S.A. and Japan) and an economic union (U.S.A. 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, these the drawn conclusions have to be considered with some caution.

Both the theoretical and the empirical literature on regional inflation rate dispersion is very limited thus far. 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. inflation rates includes Parsley and Wei (1996) and Cecchetti et al. (2002). Both of these studies use regional U.S. price data and show that shocks to relative prices are persistent. Cecchetti et al. (2002), e.g., find half-lives of PPP deviations for U.S. cities of almost nine years. They argue that these figures represent lower bound for convergence speeds of relative EMU prices. Evidence for inflation dynamics within the euro area is provided by Weber and Beck (2005) who use regional inflation rates from major EMU countries for the period before and immediately after the introduction of the euro. The two authors show that there are considerable within-distribution dynamics in European inflation rates. Furthermore, evidence is presented that inflation rate dispersion 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 U.S.A. and Japan and in the economic union between the U.S.A. and Canada. In a second step, we will examine the dynamics of overall dispersion in our three samples for the last twenty years. Our main focus will be on the question of how overall dispersion has evolved over time and how large dispersion in our samples is relative to that in the EMU (documented in Weber and Beck (2005)). In our third contribution, we will focus on the question of how the shape and the composition of regional inflation rate dispersion in our samples have evolved over time. To do so, we refer to a methodology known as distribution dynamics. 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.

We find that inflation rate dispersion is significant throughout the sample period in all three samples. However, our results show that there is significant mean reversion (β-convergence) in inflation rates in all considered samples. The evidence on σ-convergence is mixed. Observed declines in dispersion are usually associated with decreasing overall inflation level which indicates a positive relationship between mean inflation and overall inflation rate dispersion. When setting our results in relation with previously obtained evidence for EMU countries, we see that both the overall dispersion and the persistence of inflation rates is considerably lower within the U.S.A. and Japan than within the EMU. Only for the combined sample of U.S. and Canadian data, we find similar dispersion characteristics as for EMU regions. In line with findings from the 'border literature' these results suggest that frictions across European markets are significantly higher than they are, e.g., across North-American markets.

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 section 3. Section 4 examines the issue of σ-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 mentioned above, one important aspect of our analysis is the use of regional instead of national data. The merits of this approach can readily be seen from its application in the empirical growth literature where it has provided us with important insights into the nature of per-capita income dynamics across regions.[1] 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 (2005) use it to analyze the degree and the dynamics of international goods market integration. There are several reasons that make this approach 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 address questions that could not be dealt with otherwise. A comparison between within-country and cross-country goods market integration, as, e.g., done in Engel and Rogers (1996), can only be performed with this type of data. Likewise, an investigation of inflation rate dispersion in EMU, as done in Weber and Beck (2005), is difficult to perform on the basis of only twelve national CPI observations per period. 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 give us some insights into further developments within the EMU. In this context, the study of U.S. cities is probably

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

most helpful. A drawback of the use of regional data is that they are not readily available and thus have to be collected in a time-consuming 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 case 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 to a satisfactory degree.

We have compiled regional inflation data for the United States, Canada and Japan. An overview of the 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 in Japan. U.S. and Canadian CPI data are available for the time period 1980 - 2002, Japanese data are available for the period 1985 - 2000.[2] In our estimation analysis, we are looking at three different samples constructed from these data: a sample of U.S. locations, a sample of Japanese locations and a joint sample of U.S. and Canadian locations. The study of U.S. and Japanese regional inflation rate dynamics is motivated by two major aspects: First, it allows us to set results found in Weber and Beck (2005) for the EMU in relation to the evidence from well-established monetary unions (where within-country inflation rate dispersion is not seen as a problem). Secondly, the results from this analysis provide us with a benchmark for future developments of inflation rates within the EMU. The joint U.S./Canadian sample is chosen to serve as an 'upper' benchmark for the EMU. Due to the lack of a common monetary policy, we expect that the size and the persistence of regional inflation rate differentials within the EMU lie somewhere in between those of the U.S.A. and Japan on the one hand and those of the joint U.S./Canadian sample on the other hand.

All data are annual and are available in index form. Inflation rates, π_t, are computed as annual percentage changes in the price index as:

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

where π_t denotes the inflation rate in period t, and P_t represents the consumer price index in t.

To get an idea of the extent and dynamics of regional inflation rate dispersion, figures 1 to 3 plot inflation rates for all three samples. As is clear from these plots, regional inflation rate dispersion is not only of significant size but also appears to be fairly persistent. For the U.S., e.g., the difference between the lowest and the highest local inflation rate is about 3% throughout the sample period. 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 shows that all three figures share a similar time pattern. Inflation rates are highest at the beginning of the 1980s and drop until the mid 1980s. The subsequent increase until 1990/1991 is followed by a long-lasting smooth reduc-

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

tion 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 joint U.S./Canadian sample, there are some indications for a decrease in overall dispersion during the sample period. This decrease is particularly evident for the first years of the sample period and is related to the significant reduction in average inflation rates at that time. After 1985, there are no more signs of a declining overall dispersion. We will discuss this issue in more detail in section 6. One issue that cannot be addressed by pure inspection of figures 1 to 3 is whether the 'anatomy' of dispersion is changing over time, i.e., whether there is significant within-distribution dynamics. This issue is important for policy-makers since the degree of persistence in regional inflation rate differentials is of considerable importance for the question whether policy-makers have to worry about them or not.

Table 2 reports period-average inflation rates and their respective cross-regional dispersion computed for the total sample period and five-year subperiods. Looking at the five-year subperiods, it becomes clear that mean inflation rates have continuously declined in the U.S. and the U.S./Canadian sample (from an average of about 5.4% to an average of about 2.4%) in the last twenty years. The inflation pattern for Japan is somewhat different. Overall inflation has always been far below that of the U.S.A. and Canada with an maximum average inflation rate of 1.26% in the first half of the nineties and an extremely low average inflation rate of only 0.39% in the second half of the last decade. 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. (2002) 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 rate dispersion across regions.

Before, however, we briefly want to address the question of possible reasons for the observed inflation rate differentials. In figure 4 we give an overview of possible inflation determinants.[3] 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 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

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

inflation (by reducing the demand for money, e.g.), but of course actual inflation also leads to adjustments of inflation expectations.
When going over the individual factors, we can observe that they differ with respect to the regional '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. Thus, as most of the factors that might have an impact on inflation have a 'regional' dimension, we should expect inflation differences to exist at least in the short and medium-run. 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.

3 β-Convergence of Regional Inflation Rates

Figures 1 to 3 showed that there is considerable dispersion across regional inflation rates in the considered samples. However, observed inflation differentials are only troublesome, when they are highly persistent. In this paper, we are examining the degree of persistence across inflation differentials based on two methodologies: First, we examine mean-reverting behavior (β-convergence) using standard panel unit root methods and secondly, we use distribution dynamics (see section 5).
Following the soccer-league picture by Sala-i Martin (1996a), studying the presence of β-convergence corresponds to examining 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 underlying the differences in inflation rates. Since some of these factors such as indexing differences or productivity differentials, can be very 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 for our three samples (total period and five-year subperiods). When there is 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 the pictures - including the U.S./Canadian case - shows that there are clear indications of β-convergence in our samples. However,

it is not clear whether there are differences in convergence speeds across samples. The plots for the subperiods indicate that the pattern of mean reversion has been relatively stable over the last fifteen to twenty years.

To address the issue of β-convergence more formally, we refer to panel unit root methods as developed by Levin and Lin (1993).[4] 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 rate convergence is based on the following equation

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

Here, Δ denotes the one-period change of a variable and θ_t represents a common time effect. $\epsilon_{i,t}$ is assumed to be a (possibly serially correlated) stationary idiosyncratic 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. The number of included lagged differences is determined using the top-down approach suggested by Hall (1994). To take control of cross-sectional dependence, we subtract the cross-sectional mean such that equation (2) becomes

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

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

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

To see whether mean-reverting behavior in inflation rates is present, we test the null hypothesis that all ρ_is are equal to zero against the alternative hypothesis that they are all smaller than zero. If we can reject the null hypothesis of non-stationarity, 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 ρ (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 ρ are biased downward (see Nickell (1981)). To correct for this downward bias, we use the adjustment factor that Nickell suggests. Critical values for the test statistics are obtained using a parametric bootstrap based on 5,000 simulations of the data-generating process under the null hypothesis.

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

Results are presented in table 3. This table reports estimated half-lives both for the total period and ten-year 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 rate convergence is largest for the Japanese sample, followed by the U.S. and the U.S./Canadian sample. For Japan, we obtain half-lives of inflation rate deviations from the cross-regional mean of around six months. For the U.S., half-lives lie slightly above and are in the range of about one year. Interestingly, our results for the subperiods indicate that half-lives might have increased considerably in the last decade in both of these samples. As estimates for ρ indicate, most of the differences in half-lives found for the total period and subperiods are due to the adjustment factor the influence of which becomes larger the smaller the considered time period is. Interestingly, differences in estimated half-lives between the U.S./Canadian sample and the two considered monetary unions are relatively small. To a large degree, this is probably an implication of the fact that Canadian monetary policy is considerably influenced by 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 β-convergence indicate half-lives of deviations from the cross-sectional mean of less than or close to one year both in well-established monetary unions and for the 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 rate dispersion has been reached in the considered samples. Half-lives for the monetary unions show that convergence is relatively fast and poses no problems for monetary authorities. This is especially true for Japan.

Comparing the reported figures with those found for the EMU (see Weber and Beck (2005)), 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 are considered 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.A. 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 is more difficult. Depending on one's point of view these differences can be interpreted in several ways: For a critic of the euro, they reflect the troublesome heterogeneity of countries in EMU. This argument is supported by the finding that mean reversion in inflation rates of two countries that share neither a common monetary nor fiscal policy is higher than that for EMU member countries. A proponent of the euro will assess the evidence differently. First, he will point out that there is inflation rate convergence in the EMU, only the speed is relatively small. Secondly, he will - as we did above - argue that linkages between the U.S.A. and Canada are very close. Thus, the degree of heterogeneity across these countries might be of moderate size. Then, existing inflation rate dispersions across the U.S.A.

and Canada do not necessarily represent unsustainable levels of dispersion. And last but not least, he will mention that the EMU is still in a transition phase. In this sense, particularly results for the U.S.A. represent a 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 σ-Convergence of Regional Inflation Rates

In the previous section, we examined the degree of β-convergence in regional inflation rates. In the growth literature, another concept of convergence, denoted as σ-convergence, has been extensively studied.[5] This concept focuses on the evolution of the overall dispersion of a variable over time. Applied to our case, this means to examine whether overall regional inflation rate dispersion has decreased, has remained constant or has increased throughout the sample period. For a central banker, this question is of great importance. When overall dispersion is large, he faces the difficult task to meet conflicting demands from different regions: Whereas for low-inflation areas an expansive 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 rate dispersion for the U.S., Japan and the U.S./Canadian sample. Additionally, we will relate the findings to previously obtained results for the euro area.

Following the empirical growth literature, we examine σ-convergence by looking at the evolution of the standard deviation of regional inflation rates over time. 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 (around 0.4) is considerably smaller than that of the U.S. (0.6 and higher) or the U.S./Canadian 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 the included countries. As the literature on national and international goods market integration shows[6], 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.A. than for Japan. The second difference between the North-American samples and the Japanese sample concerns the dynamics of overall dispersion. For the North-American samples, we find signs of decreasing overall dispersion. This is particularly pronounced for the deflation phase at the beginning of the 1980s, but even after this period there is a further, however much smoother decline until the mid 1990s. For the second half of the 1990s, overall dispersion has slightly increased. For Japanese

[5] The terms β- and σ-convergence date back to the Ph.D. thesis of Sala-i-Martin, see Sala-i Martin (1990).

[6] See, e.g., Engel and Rogers (1996) and Parsley and Wei (1996) for North-American and Beck and Weber (2005) for European evidence.

prefectures, overall dispersion does not show a comparable decline. On the contrary, it even slightly increased in the sample period. The interesting coincidence of movements between the mean of inflation rates and their dispersion will be explored in more detail in section 6.

Unlike in the case of β-convergence, the differences in overall dispersion between the U.S. (U.S./Canadian) sample and the EMU are not so pronounced. Weber and Beck (2005) show that there has been a significant decline in EMU regional inflation rate dispersion from about 1.5 to about 0.7/0.8 in the 1990s. These numbers 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 of the same extent as it is for the U.S.A. This result is supportive of the upper mentioned view of euro proponents. The smaller degree of mean reversion may then 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 β-convergence on one side and missing or 'negative' evidence of σ-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 β- and σ-convergence. This can best be done by referring to Sala-i Martin (1996b) who shows that 'β-convergence is a necessary condition for σ-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. This figure represents the different possible combinations of β- and σ-convergence. In panel (a) of figure 14, β-convergence induces σ-convergence. On the other hand, as panel (b) shows, the absence of β-convergence implies the absence of σ-convergence. The most interesting case is demonstrated in panel (c). Here, β-convergence occurs, however, σ-convergence cannot be observed. In other words, the strong evidence of β-convergence found in the last section does not imply σ-convergence. In fact, it is - as we have seen - even possible to find increasing dispersion even though β-convergence is present. This happens 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 are 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 distribution dynamics methodology that is leant from the empirical growth literature where is has been employed to study the composition of worldwide income distribution over time.[7] In this paper it will used to examine the

[7] 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).

evolution of the composition of regional inflation rate dispersion over time. Knowing how the composition of the tails of existing inflation rate 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 (2005), we perform a similar exercise using European regional inflation rates. We find that there is significant within-distribution dynamics: Large deviations of inflation rates from the cross-regional mean are expected to decrease considerably at a one-year horizon. Transition probabilities of staying at the left or right tail of the distribution of mean-inflation rate deviations are estimated to be between 50% and 60%. Comparing these results with analogous findings for the two monetary unions considered in this paper, can help to obtain a better assessment of the prevailing dynamics in Europe. The U.S./Canadian case on the other hand again serves as an upper benchmark for the EMU.

The idea behind distribution dynamics is to find a law of motion that describes the evolution of an entire distribution over time. Following the growth literature, we use a Markov process to describe the dynamics of the distribution of cross-regional mean-inflation rate deviations, denoted as F_t. More specifically, the dynamics of F_t is modelled as an AR(1) process in the following way:[8]

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

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, denoted as X_t, this operator is either interpreted as the transition function/stochastic 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

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

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, that any discretization will be arbitrary. Insofar,

[8]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).

the figures presented in this section have to be treated with a caveat but are very useful for practical considerations.[9] 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} (with $i, j = 1, 2, \ldots, n$) denotes the conditional probability of the event that a region's inflation rate that is currently in inflation state i will move to state j in the next period.
Figures 15 to 20 present the results for the continuous case. For each sample, the three-dimensional graph represents the surface plot of the estimated stochastic kernel for regional mean-inflation rate deviations. On the x-axis (denoted by t), we plot period's t inflation rate deviations from the cross-regional mean and on the y-axis (denoted by $t + 1$), we plot period's $t + 1$ inflation rate deviations from the cross-regional mean. On the z-axis, we plot the transition density function $p(x, y)$ associated with the stochastic kernel $P(x, A)$.[10]. If the probability mass was concentrated along the diagonal of the x-y plain, then any existing deviation 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 x-axis - then period-t deviations would be basically expected to vanish until next period. The lines of the contour plots of the estimated stochastic kernels (left panel of the figure below the surface plots) represent lines of constant density of the respective surface plots. Conditional expected mean-inflation rate deviation for the period following the observation period are plotted to the right of the contour plots. Figures 15 to 20 show a pattern of within-distribution dynamics that is compatible with the evidence that we obtained in the section on β-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 an inflation rate that is relatively large or small compared to the cross-regional mean-inflation rate to have a similarly large or small relative inflation rate one year later is relatively small. On the other, there is a large likelihood that the region's relative inflation rate 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 mean-inflation rate deviation (right panel of figure 16). Next period's expected mean-inflation rate deviation is, e.g., only around

[9] 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) show that a regenerative discretization instead of our 'naive' discretization would probably not change our main results dramatically.

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

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

with y denoting elements in A When A is identical to the underlying state-space (R), the transition density function integrates to one

-0.5% given it is -2% in this period. Thus, a considerable movement towards the cross-regional mean is expected. However, as the contour plot (left panel of figure 16) shows, there is also a non-negligible probability that an existing inflation rate 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 conditional mean deviation (right panel of figure 18). The contour plot (left panel of figure 18) shows that the probability of a large mean-inflation rate deviation 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, 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 findings for the joint U.S./Canadian sample, results confirm the smaller degree of dynamics towards the mean found in the section on β-convergence. Although current deviations from the cross-regional mean are expected to decline, the extent in which this happens is significantly smaller than that we observed, e.g., for the U.S. sample. This result is 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 of our three samples to those that we obtained in Weber and Beck (2005) for the EMU area. The closest 'match' with the results for EMU inflation rate behavior is again found for the U.S./Canadian sample. The contour plots and the estimates for conditional expected changes in mean-inflation rate deviations between these two samples are similar. Thus, inflation rate dynamics within the EMU is more similar to that of the joint U.S./Canadian sample than it is relative to the U.S. sample, although the U.S.A. and Canada do not share a common monetary policy. As we argued above, this result probably reflects two features of the euro area that have often been cited in the past by euro critics. The first is the absence of a central fiscal authority that can use transfer payments 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 persistent effects of asymmetric shocks which is reflected in the weaker dynamics towards mean reversion in our inflation rate data.

While an examination of the continuous state-space is more appropriate for inflation rates than a corresponding discrete-state space analysis, 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 will provide transition probabilities for our three samples in the following. To do so we start by dividing the continuous state-space into five discrete states. Boundaries for the states 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 conclusions that we have drawn from the graphical findings for the continuous case. For the U.S., the probability that a region's inflation rates deviates by more than 0.7% from the cross-regional mean for two subsequent periods is about 50% (42% for negative deviations and 57% for positive deviations). As the first and the fifth row of the upper panel of table 4 show, 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 towards the mean is more pronounced for Japanese prefectures. From the numbers chosen to classify states, the lower degree of overall dispersion becomes evident. Looking at the second panel of table 4, we can see that the probability of remaining in an 'extreme' state (characterized by large mean-inflation rate deviations) is only about 30%. Thus, the likelihood of moving closer to the mean is 70%. Moreover, there is a considerable probability that existing mean-inflation rate deviations 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 0.30% from the mean-inflation rate next period when current mean-inflation rate deviation is below 0.10% is 30%. Overall, the results for Japan show that there is large within-distribution dynamics across prefectures.

As for the continuous case, the distribution patterns for the joint U.S./Canadian case are most similar to those observed for regional EMU inflation rates. As the third panel of table 4 shows, the transition probability for the U.S./Canadian economic union has a structure comparable to that of the EMU (see table 8 of Weber and Beck (2005)). Table entries show, that the tendency of regions in the left or right tail of the distribution to exhibit small shifts towards the mean is as pronounced as for the U.S.A. However, large shifts are less likely to occur.

In the next section, we want to perform a little exercise that will provide us with 'critical values' for mean inflation rates below which a significant share of regions faces negative inflation rates. These values can be directly used by U.S. and Japanese policy-makers as indicators for significant deflationary threats in their respective countries. They additionally have some benchmark character for the EMU.

6 Mean Inflation Rates and Inflation Rate Dispersion

In recent months, there has been a vivid discussion on how large the probability is that countries like the U.S.A. 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'. Ahearne et al. (2002) take the Japanese deflation experience of the 1990s to draw lessons for monetary policy in countries that are 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 in 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 zero. When dispersion is large, this can occur to a large extent. Then, debt holders in these regions may suffer from ever-increasing real values for their debts which might get local banks into trouble. Our strategy in examining this issue is as follows: We will start by using the regional dimension of our data to approximate existing empirical inflation rate dispersions by a theoretical counterpart. This enables us to compute 'critical values' for country-wide average inflation rates. These 'critical values' indicate the portion of regions that face negative inflation rates at a given national average inflation rate.

Before proceeding with searching for an adequate theoretical distribution that fits the main characteristics of our data, we shortly want to turn our attention to an observation that we already mentioned above and that becomes of interest in this section. The plots of the regional inflation rates (figures 1 to 3) together with the plots of the evolution of the dispersion of regional inflation rates over time (figures 11 to 13) indicate a positive relationship between the overall level of inflation rates and the degree of regional dispersion. This is particularly pronounced for the early period in the U.S. sample where the strong reduction in inflation rates is accompanied by a strong reduction in overall dispersion. When computing 'critical values', we will take this relationship into account.

Similar to our finding of a positive relationship between a country's average inflation rate and its regional inflation rate dispersion, a large branch of literature has empirically examined an analogous relationship between a country's inflation rate and its cross-sectional dispersion.[11] 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 has also 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. 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 rate dispersion across re-

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

gions.
In search for an appropriate theoretical distribution necessary to compute 'critical values' we particularly need to make sure that the left and right tails of the theoretical distribution capture their empirical counterparts sufficiently well. Fortunately, normal distributions seem to fit current inflation rate dispersions sufficiently well for our purposes. Figure 21 plots empirical density estimates of regional inflation rate dispersions in 2000 for all three samples. Additionally, the plots contain a plot of the normal density function that we use to approximate the respective empirical distribution. Although the fit is naturally not perfect, we think that it is good enough for the purpose of computing 'critical' mean inflation rate values. Given our choice of using normal distributions, we only need to compute first and second moments of the observed empirical regional inflation rate 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.[12] Secondly, we have to make sure that the supposed relationship between the average level of inflation rates and its cross-regional dispersion is taken into account. We do this by estimating a functional relationship between the observed mean inflation rate 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 μ_t denotes the sample-wide average inflation rate in period t and σ_t denotes the standard deviation of regional inflation rates in period t. The use of absolute values for the mean inflation rate shows that we assume a symmetric relationship for positive and negative mean inflation rates. Although one can certainly doubt this assumption for a variety of reasons, we think that it can serve as a good working hypothesis. Estimation results are presented in table 5. Both for the U.S. and the joint U.S./Canadian sample a significant relationship between the mean inflation rate and its regional dispersion exists. 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 are 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 different perspectives. First, we calculate 'critical values' for the sample-mean of inflation rates at which 1%, 2.5%, 5%, 10% and 25% of all regions face negative inflation rates.

[12] 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.

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 and p_{crit} is the proportion of regions facing inflation rates below zero. μ_{crit} denotes the value of the sample-mean inflation rate at which p_{crit} percent of all regions face negative inflation rates. The expression $\sigma(\mu)$ reflects the above stated relationship between the mean of regional inflation rates and their dispersion. To determine 'critical values' for μ, we set π equal to zero and solve the equation for μ_{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 (2005)). The figures are very illustrative. Given the current degree of inflation rate dispersion, 5% of all regions in the U.S.A. will have inflation rates below zero when the nation-wide inflation rate is about 1%. When national average inflation drops to 0.79%, the share of regions with inflation rates below zero rises to 10%. Thus, only for relatively low nation-wide inflation rates the share of regions with negative inflation rates becomes notable. Can we therefore totally ignore inflation rate 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 about 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. Evidence for Japanese prefectures reflects our previous findings that regional dispersion in Japan is of minor size. Only when the country-wide inflation rate is 0.5%, 10% of all regions have negative inflation rates. As in the last section, the results that are closest to those of the EMU are those for the joint U.S./Canadian sample. As dispersion is larger for this sample than for the other two considered samples, 'critical values' are higher. However, they are still considerably smaller than those for the EMU despite the U.S.A. and Canada do not share a common monetary policy.

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 rates. It is supposed to give us 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 sample-wide inflation rates fall below this point, the share of regions with negative inflation rates rises fast. Given the supposed overestimation of actual inflation by reported numbers, a central bank should definitely not consider pushing inflation below 1%. Interestingly, this finding is applicable for both the U.S.A. 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 regional inflation rate dispersions 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 low-inflation regions, a restrictionary policy is necessary for high-inflation regions. Thus, the question of inflation rate 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 rate dispersions in the considered samples are and how they evolve over time. Our data for the U.S., Canada and Japan show that inflation rate dispersions are considerable in all considered samples. 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 β-convergence we found significant mean reversion in inflation rates. The lowest half-lives of around six months are found for Japan, for the U.S.A. we document half-lives of around nine months and for the U.S./Canadian sample half-lives of more than one year are found. Although these findings indicate some persistence in the factors causing inflation rate dispersion, half-lives reflect considerable dynamics towards the mean. Looking at the evolution of overall inflation rate dispersions, we found strongly declining dispersion in the early 1980s for the U.S. and the joint U.S./Canadian sample and no or only minor declines afterwards. As there are strong indications of considerable mean reversion in regional inflation rates in all samples, the prevailing dispersions in inflation rates do not pose major problems to monetary authorities in the U.S.A. or Japan. Such a conclusion is supported by our results in the section on distribution dynamics. Modelling the evolution of the distribution of regional inflation rates as an AR(1) process, we find indications of dynamics towards the cross-sectional mean in all samples. Again, the largest dynamics are found for Japan, followed by the U.S. and the joint U.S./Canadian sample. Estimated transition probability matrices indicate that there is a 50%-probability that large mean-inflation rate deviations reduce considerably within one year.

In an attempt to assess the significance of existing dispersions of regional inflation rates for the ongoing discussion on deflationary threats, we are providing 'critical values' for aggregate inflation rates. In computing these values for the U.S. and the combined U.S./Canadian data, we are using a statistically significant relationship between the prevailing nation-wide mean inflation rate and its dispersion. 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 underestimation of true inflation by reported number, we thus support Bernanke's buffer zone argument to prevent deflation and suggest a size of at least 1%.

One of our most striking results concerns the comparison of the findings in this paper with the evidence for EMU countries (see Weber and Beck (2005)). 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 the EMU do. Depending on one's point of view, these findings can be interpreted differently. Even proponents of the new currency have to admit that there are obviously 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 countries is not adequate does not necessarily has to be drawn for several reasons. First, linkages between the U.S.A. and Canada are large with respect to many dimensions. Not only does a free trade arrangement between these countries exist but there is also 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 is not one of two economic entities that are subject to large asymmetric dynamics. Secondly, it is very likely that there will be further steps towards convergence across EMU member countries in the next few years. Thirdly, even if there will be no more process towards further integration, an important question is whether the current extent of regional inflation rate dispersion in the EMU is unsustainable. Due to a lack of comparable events in the past, this question cannot be ultimately answered here. However, comparing results for the EMU and the U.S.A., we see that the overall level of regional inflation rate dispersion is not much larger in the EMU than it is in the U.S.A. Additionally, there are significant tendencies towards the mean in both samples. The difference only concerns the speed at which convergence occurs. As these differences very likely reflect larger segmentations in markets across EMU member countries, the ECB is certainly right when it asks politicians to remove existing rigidities across European markets.

8 Tables

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

U.S.A. (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 **Coverage**: 1980 - 2004
Canada (12 provinces)
Prince Edwards Islands, Alberta, New Brunswick, Nova Scotia, Quebec, Saskatchewan, New Foundland, Ontario, British Colombia, Yukon, Manitoba, Yellowknife
Source: Statistics Canada **Coverage** 1980 - 2004 **Notes**: 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 **Coverage** 1985 - 2000

Table 2: Descriptive Statistics

U.S.A					
	1981-2004	1981-1985	1986-1990	1991-1995	1996-2004
mean	3.42	5.36	3.69	3.07	2.42
std.dvt.	0.08	1.83	2.52	0.86	0.41
Japan					
	1986-2000	1986-1990	1991-1995	1996-2000	
mean	0.96	1.23	1.26	0.39	
std.dvt.	0.05	0.60	0.29	0.45	
USA/Canada					
	1981-2004	1981-1985	1986-1990	1991-1995	1996-2004
mean	3.45	5.84	3.76	2.83	2.25
std.dvt.	0.08	5.69	3.01	1.88	0.77

Notes:

1) The mean inflation rate (mean) is computed as the cross-sectional mean of all regional mean inflation rates (geometric mean) included in the respective sample. The computation of the standard deviation is likewise based on the cross-section of the geometric means of all regional mean inflation rates included in the respective sample.

2) Standard deviations are multiplied by 10,000.

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

Sample	ρ	ρ_{adj}	$t - stat$	p-value	half-life	half-life (adj.)
			U.S.A.			
1981-2004	0.42	0.48	-16.64	0.000	0.8	1.0
1981-1990	0.33	0.49	-12.33	0.000	0.6	1.0
1991-2004	0.55	0.68	-8.54	0.000	1.1	1.8
			Japan			
1986-2000	0.28	0.37	-18.53	0.000	0.5	0.7
1991-2000	0.26	0.41	-16.00	0.000	0.5	0.8
			USA/Canada			
1983-2004	0.39	0.45	-20.10	0.000	0.7	0.9
1983-1990	0.56	0.85	-9.87	0.006	1.2	4.2
1991-2004	0.57	0.7	-10.42	0.000	1.2	2.0

Notes:
1) Results are based on the equation:

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

where $\tilde{\pi}_{i,t}$ denotes the deviation of region's i inflation rate from the cross-regional mean. A more detailed description of the estimation procedure is given in section A.
2) Bias adjustment is done using the formula given by Nickell (1981).

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

Transition Probabilities for the U.S.A.					
Dev. in t	Dev. in $t+1$				
	< -0.7	-0.2	0.2	0.7	> 0.7
< -0.7	0.42	0.24	0.14	0.14	0.07
-0.2	0.24	0.28	0.24	0.17	0.07
0.2	0.18	0.32	0.2	0.21	0.1
0.7	0.05	0.19	0.3	0.33	0.12
> 0.7	0.02	0.09	0.1	0.22	0.57
Transition Probabilities for Japan					
Dev. in t	Dev. in $t+1$				
	< -0.30	-0.10	0.1	0.3	> 0.30
< -0.30	0.28	0.27	0.26	0.12	0.07
-0.10	0.19	0.28	0.21	0.2	0.13
0.10	0.19	0.21	0.25	0.23	0.11
0.30	0.13	0.17	0.22	0.28	0.19
> 0.30	0.07	0.22	0.19	0.21	0.32
Transition Probabilities for USA/Canada					
Dev. in t	Dev. in $t+1$				
	< -0.7	-0.2	0.2	0.7	> 0.7
< -0.7	0.46	0.24	0.12	0.1	0.08
-0.2	0.24	0.26	0.26	0.14	0.1
0.2	0.21	0.24	0.19	0.19	0.17
0.7	0.09	0.15	0.26	0.2	0.31
> 0.7	0.07	0.1	0.17	0.18	0.49

Notes:

1) Table entries report conditional probabilities for the event that an observation which is in period t in the state indicated in column one moves to one of the states indicated in columns two to six in period $t+1$. The variable under consideration is the deviation of a certain region's inflation rate from the cross-sectional mean of inflation rates. Each state includes all inflation rate deviations that lie within the indicated range. The state -0.20, e.g., comprises all inflation rate deviations that lie in the range $[-0.70, -0.20[$. States were chosen such that each state has approximately the same number of observations.

Table 5: Examining the Relationship between the Cross-Regional Mean of Inflation Rates and Their Dispersion

Estimated Equation: $\sigma_t = \alpha + \beta\mu_t + \epsilon_t$			
α	β	R^2_{adj}	$s.e.r.$
	U.S.A		
0.006	0.062	0.18	0.002
0.001	0.029		
	Japan		
0.004	-0.015	0.05	0.001
0.0003	0.019		
	USA/Canada		
0.005	0.093	0.25	0.003
0.001	0.036		

Notes:

1) σ_t denotes the standard deviation of regional inflation rates in period t, μ_t denotes their mean.

2) Numbers in brackets denote standard deviations of the estimated coefficients.

3) The term R^2_{adj} denotes the adjusted coefficient of determination, the term $s.e.r.$ denotes the standard error of the regression.

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

'Critical' Average Inflation Rates				
Prop. of 'Defl.' Regions	U.S.A.	Japan	U.S.A./Canada	EMU
1%	1.53	0.88	1.64	1.99
2.5%	1.26	0.75	1.32	1.53
5%	1.04	0.63	1.07	1.20
10%	0.79	0.49	0.80	0.87
25%	0.40	0.26	0.40	0.41
Mean Inflation Rate and Percentage of Regions with Deflation				
Mean Infl. Rate	U.S.A.	Japan	U.S.A./Canada	EMU
2.00	0.18	0.00	0.33	0.98
1.90	0.26	0.00	0.45	1.19
1.80	0.38	0.00	0.61	1.45
1.70	0.55	0.00	0.82	1.78
1.60	0.79	0.00	1.11	2.18
1.50	1.12	0.00	1.49	2.68
1.40	1.58	0.01	1.99	3.30
1.30	2.20	0.02	2.65	4.07
1.20	3.03	0.07	3.51	5.01
1.10	4.12	0.17	4.61	6.17
1.00	5.55	0.40	6.02	7.61
0.90	7.38	0.88	7.81	9.36
0.80	9.68	1.78	10.04	11.51
0.70	12.54	3.35	12.80	14.11
0.60	16.00	5.89	16.14	17.24
0.50	20.13	9.71	20.14	20.97
0.40	24.94	15.05	24.83	25.37
0.30	30.42	21.98	30.22	30.48
0.20	36.50	30.39	36.28	36.33
0.10	43.08	39.91	42.92	42.86
0.00	50.00	50.00	50.00	50.00

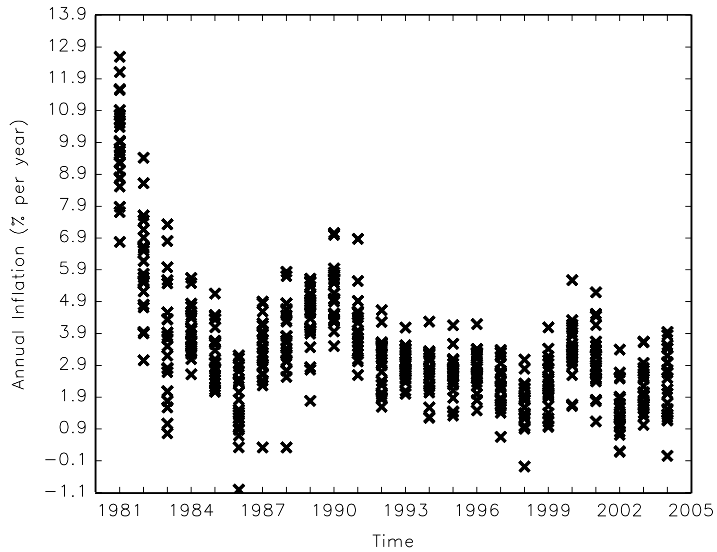
Notes:
1) Mean inflation rates (Mean Infl. Rate) are computed by weighting each regional inflation rate, $\pi_{i,t}$, with the respective region's share in total GDP, i.e.,

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

γ_i represents the share of region's i GDP (denoted as GDP_i) in total GDP (given by the sum over all GDP_i). γ_i is thus computed as $\gamma_i = \frac{GDP_i}{\sum_{i=1}^{N} GDP_i}$.

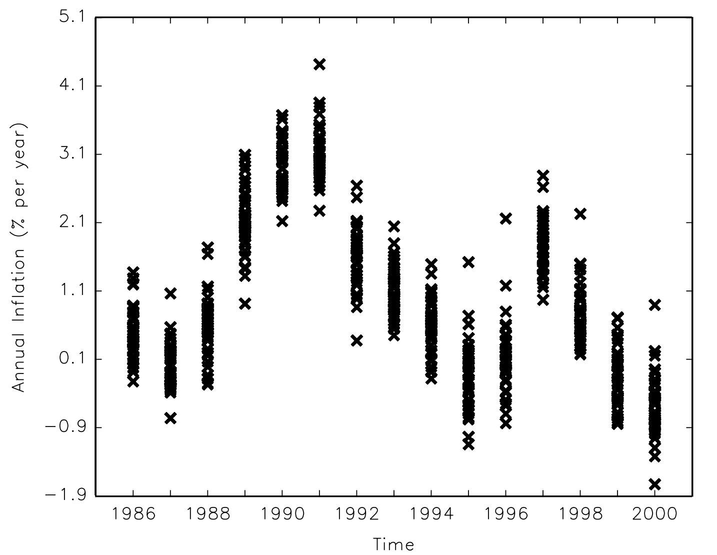
9 Figures

Figure 1: Inflation Rates of U.S. Metropolitan Areas



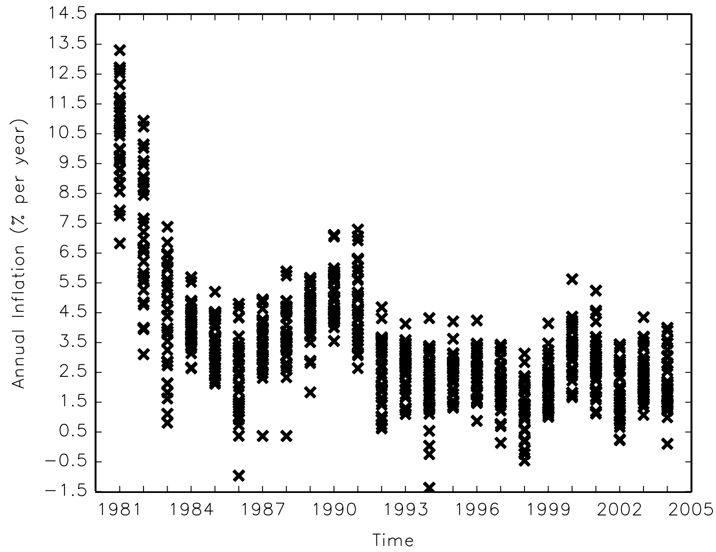
Note: *Figure 1 plots annual inflation rates of U.S. metropolitan areas. Inflation rates are computed as annual percentage changes in the underlying price index.*

Figure 2: Inflation Rates of Japanese Prefectures



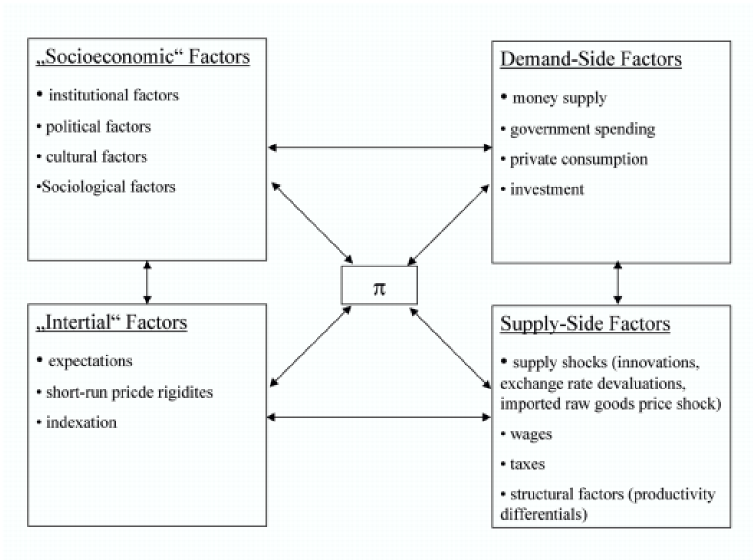
Note: *Figure 2 plots annual inflation rates of Japanese prefectures. Inflation rates are computed as annual percentage changes in the underlying price index.*

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



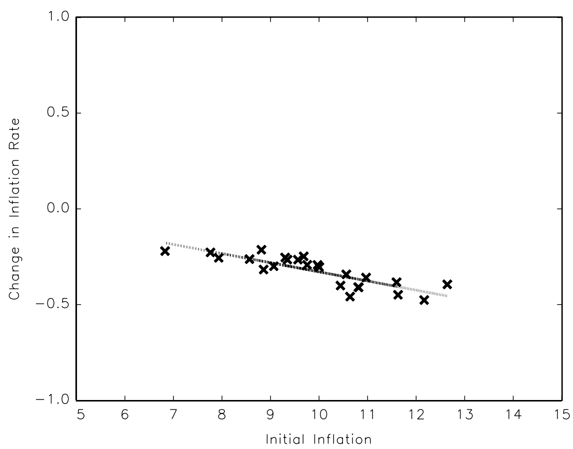
Note: *Figure 3 plots annual inflation rates of U.S. metropolitan areas and Canadian provinces. Inflation rates are computed as annual percentage changes in the underlying price index.*

Figure 4: Inflation Determinants: Overview



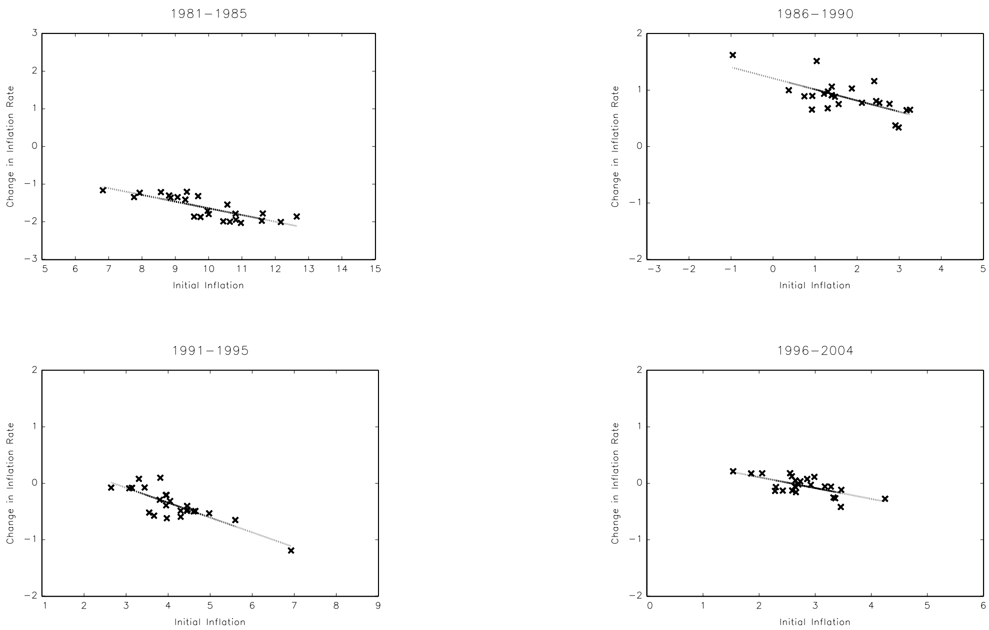
Note: *Figure 4 gives an overview of possible inflation determinants discussed in the literature on inflation. The individual factors are grouped into four categories. As the arrows indicate, individual factors are assumed to be interdependent.*

Figure 5: Change in Inflation Rates vs. Initial Inflation Rates: U.S.A., Total Period



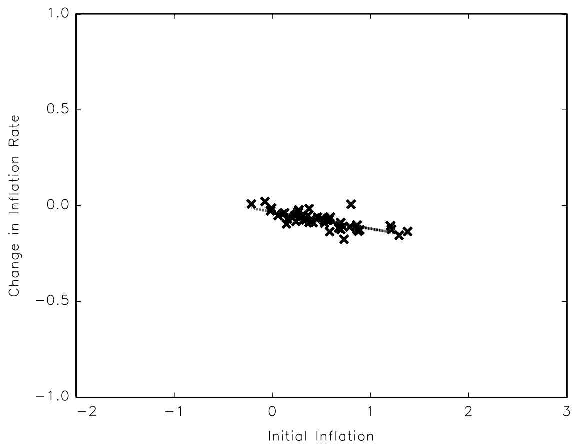
Note: *Figure 5 plots average annual changes in inflation rates between 1981 and 2002 for U.S. metropolitan areas versus their initial inflation rates in 1981. Inflation rates are computed as annual percentage changes in the underlying price index. The dotted line plots fitted values from a OLS regression.*

Figure 6: Change in Inflation Rates vs. Initial Inflation Rates: U.S.A., Subperiods



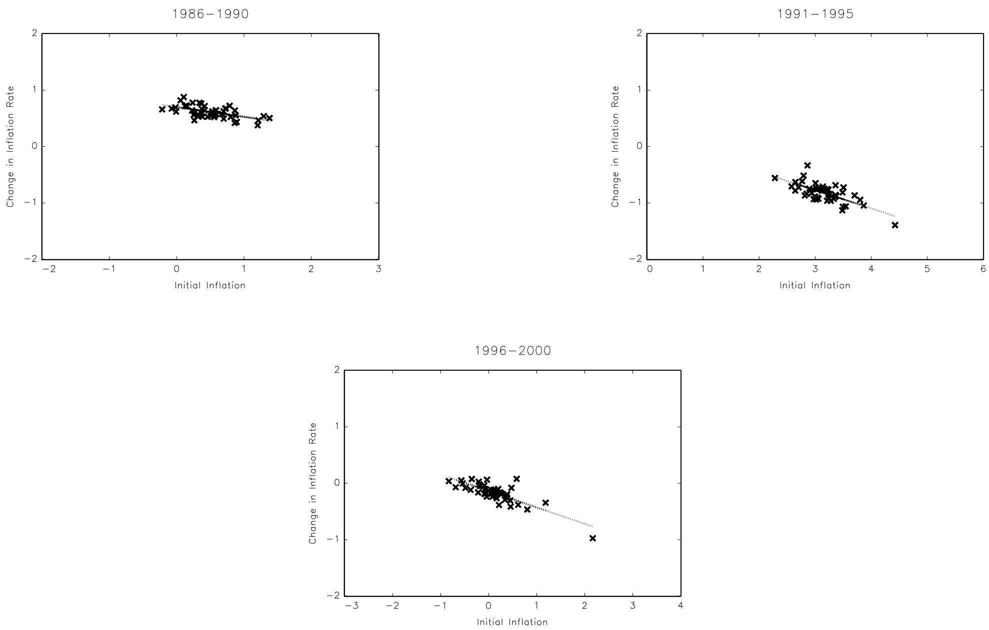
Note: *Figure 6 plots average annual changes in inflation rates for U.S. metropolitan areas versus their initial inflation rates for four subperiods. Inflation rates are computed as annual percentage changes in the underlying price index. The dotted line plots fitted values from a OLS regression.*

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



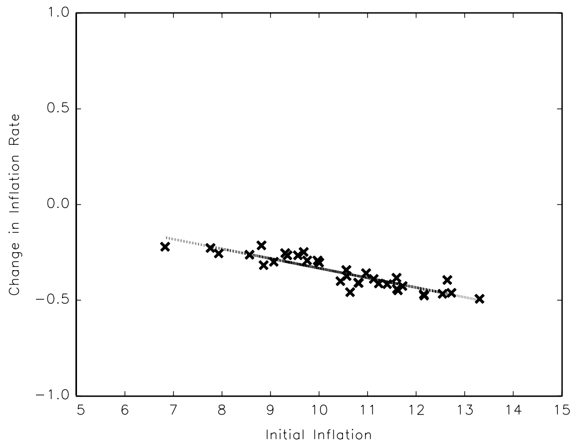
Note: *Figure 7 plots average annual changes in inflation rates between 1986 and 2000 for Japanese prefectures versus their initial inflation rates in 1986. Inflation rates are computed as annual percentage changes in the underlying price index. The dotted line plots fitted values from a OLS regression.*

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



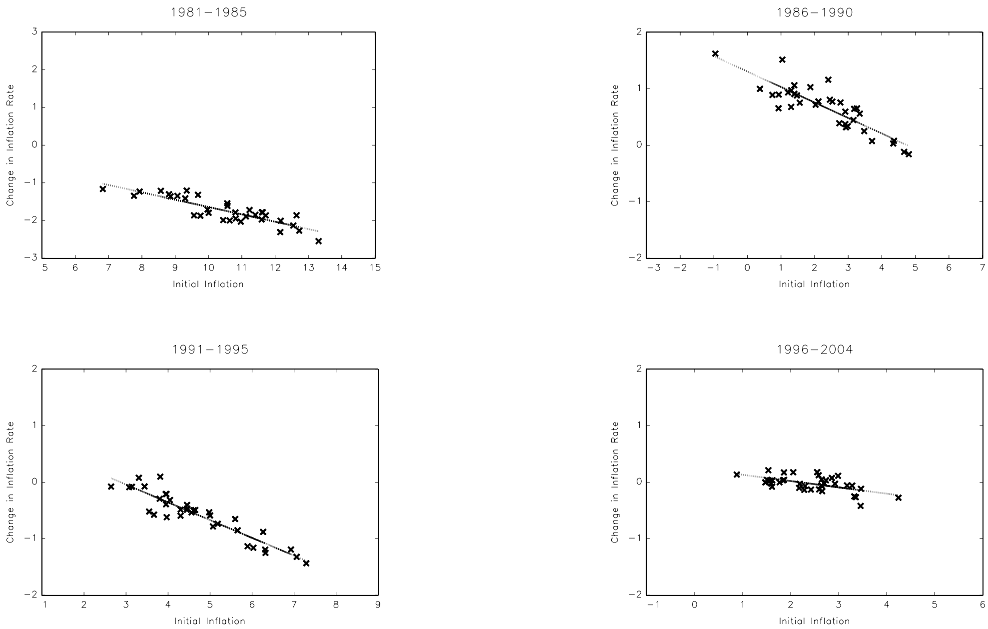
Note: *Figure 8 plots average annual changes in inflation rates for Japanese prefectures versus their initial inflation rates for three subperiods. Inflation rates are computed as annual percentage changes in the underlying price index. The dotted line plots fitted values from a OLS regression.*

Figure 9: Change in Inflation Rates vs. Initial Inflation Rates: U.S.A. and Canada, Total Period



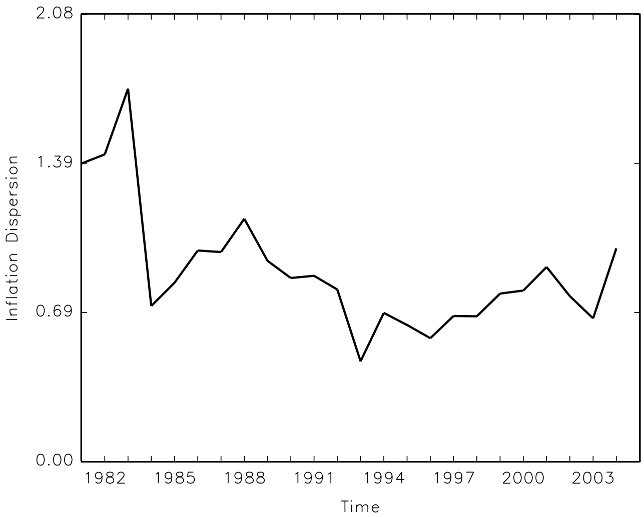
Note: *Figure 9 plots average annual changes in inflation rates between 1981 and 2002 for U.S. metropolitan areas and Canadian provinces versus their initial inflation rates in 1981. Inflation rates are computed as annual percentage changes in the underlying price index. The dotted line plots fitted values from a OLS regression.*

Figure 10: Change in Inflation Rates vs. Initial Inflation Rates: U.S.A. and Canada, Subperiods



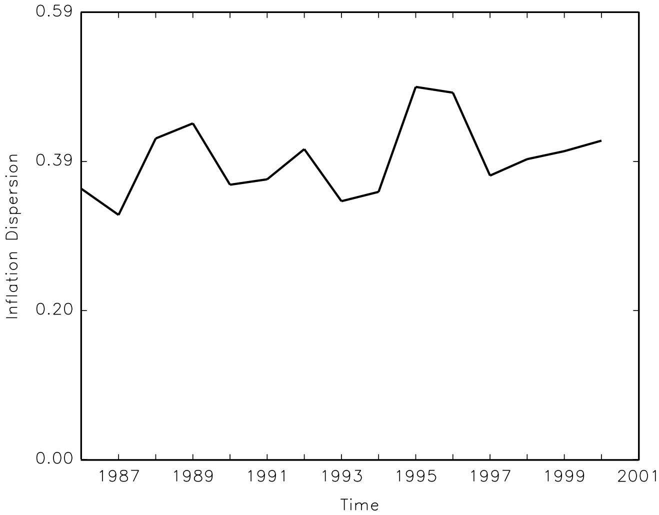
Note: *Figure 10 plots average annual changes in inflation rates for U.S. metropolitan areas and Canadian provinces versus their initial inflation rates for four subperiods. Inflation rates are computed as annual percentage changes in the underlying price index. The dotted line plots fitted values from a OLS regression.*

Figure 11: Cross-Regional Inflation Rate Dispersion: U.S.A., Total Period



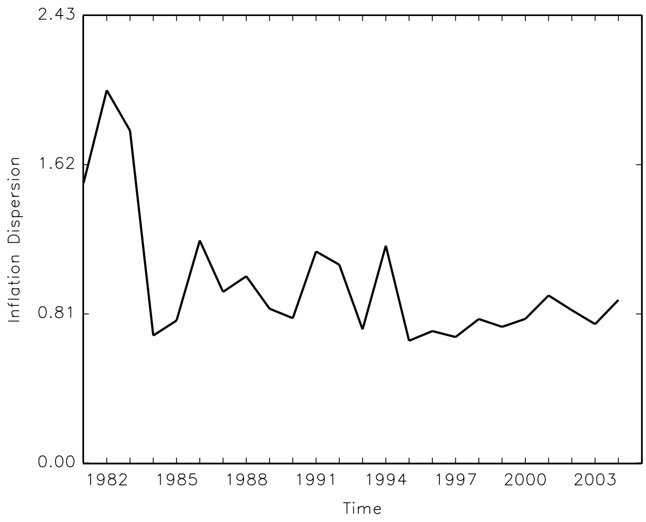
Note: *Figure 11 plots the standard deviation of U.S. regional inflation rates. Inflation rates are computed as annual percentage changes in the underlying price index. All figures are multiplied by 100.*

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



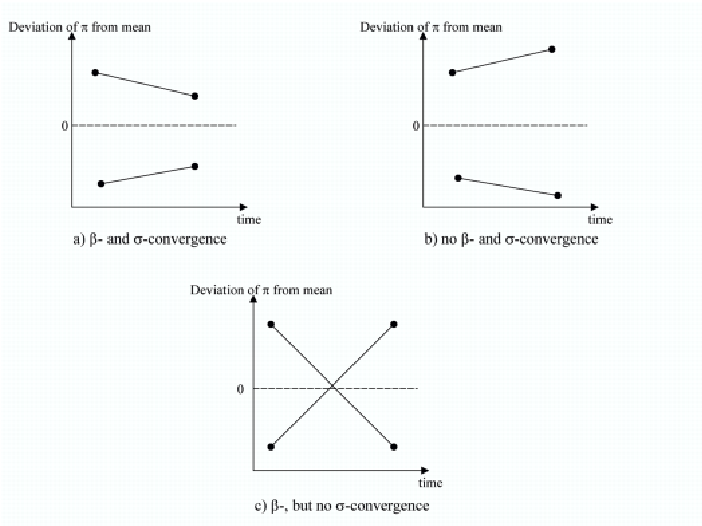
Note: *Figure 12 plots the standard deviation of Japanese regional inflation. Inflation rates are computed as annual percentage changes in the underlying price index. All figures are multiplied by 100.*

Figure 13: Cross-Regional Inflation Rate Dispersion: U.S.A. and Canada, Total Period



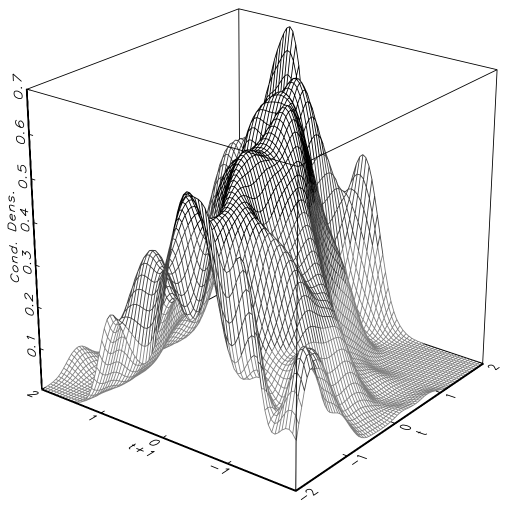
Note: *Figure 13 plots the standard deviation of U.S. and Canadian regional inflation rates. Inflation rates are computed as annual percentage changes in the underlying price index. All figures are multiplied by 100.*

Figure 14: The Relationship between β- and σ-Convergence



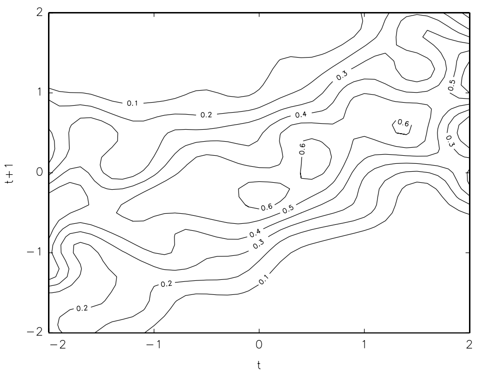
Note: *Figure 14 illustrates the relationship between β- and σ-convergence. The individual panels reflect three different possibilities of how inflation rates of two different regions can evolve over time. The graph is borrowed from Sala-i Martin (1996a).*

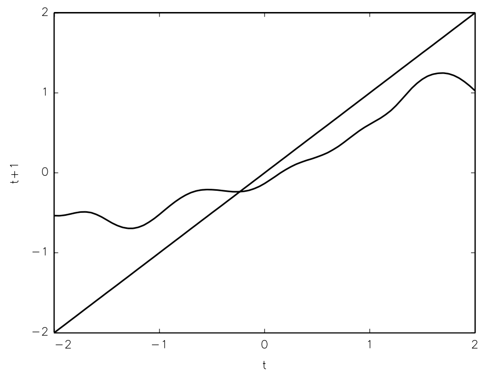
Figure 15: Surface Plot of the Estimated Stochastic Kernel for U.S. Regional Mean-Inflation Rate Deviations



Note: *Figure 15 represents the surface plot of the estimated stochastic kernel for cross-sectional mean inflation rate deviations of U.S. metropolitan areas over the period 1983 to 2002. On the x-axis (denoted by t), period's t inflation rate deviations from the cross-regional mean and on the y-axis (denoted by* $t + 1$*), period's* $t + 1$ *inflation rate 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.*

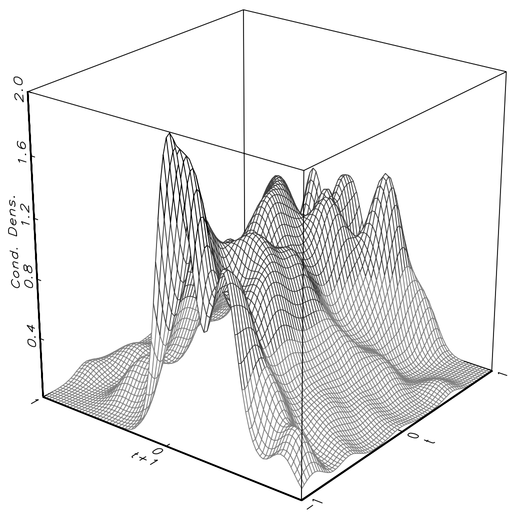
Figure 16: Contour Plot of the Estimated Stochastic Kernel and Conditional Expected Next Period's Mean for U.S. Regional Mean-Inflation Rate Deviations





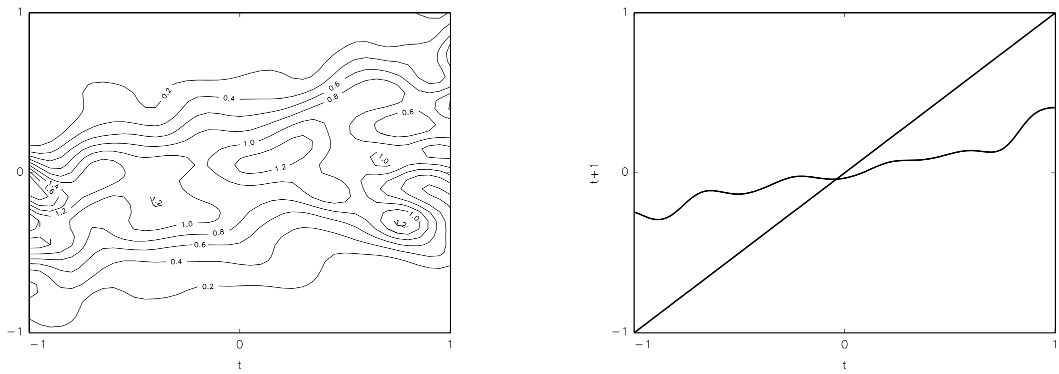
Note: *The left panel of figure 16 represents the contour plot of the transition density function* $p(x, y)$ *associated with the stochastic kernel* $P(x, A)$ *that we computed for U.S. metropolitan areas (see figure 15). The right panel of figure 16 plots expected period's* $t+1$ *mean-inflation rate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

Figure 17: Surface Plot of the Estimated Stochastic Kernel for Japanese Regional Mean-Inflation Rate Deviations



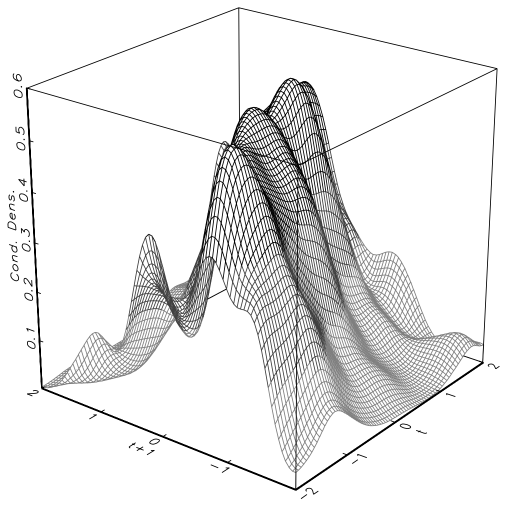
Note: *Figure 17 represents the surface plot of the estimated stochastic kernel for cross-sectional mean inflation rate deviations of Japanese prefectures over the period 1986 to 2000. On the x-axis (denoted by t), period's t inflation deviations from the cross-regional mean and on the y-axis (denoted by t + 1), period's t + 1 inflation rate 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: Contour Plot of the Estimated Stochastic Kernel and Conditional Expected Next Period's Mean for Japanese Regional Mean-Inflation Rate Deviations



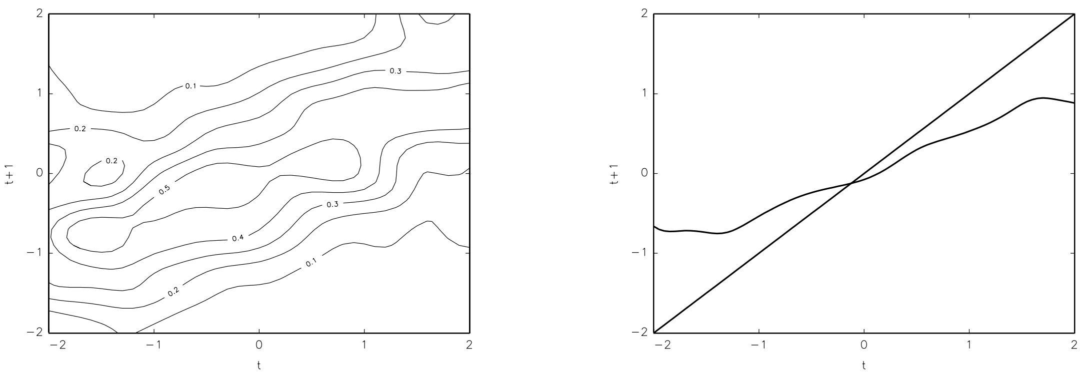
Note: *The left panel of figure 18 represents the contour plot of the transition density function $p(x, y)$ associated with the stochastic kernel $P(x, A)$ that we computed for Japanese prefectures (see figure 17). The right panel of figure 18 plots expected period's $t+1$ mean-inflation rate deviations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

Figure 19: Surface Plot of the Estimated Stochastic Kernel for U.S. and Canadian Regional Mean-Inflation Rate Deviations



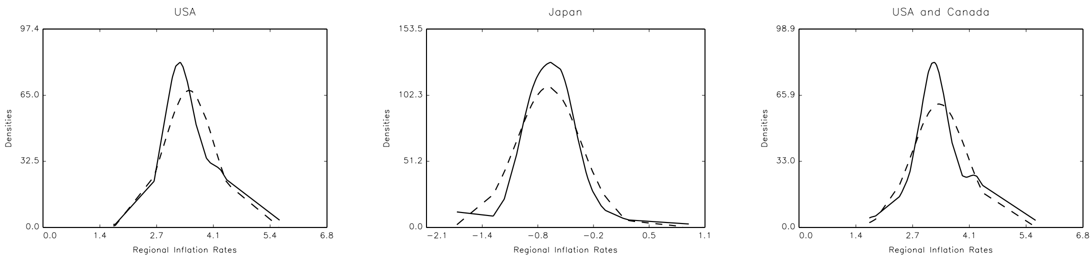
Note: *Figure 19 represents the surface plot of the estimated stochastic kernel for cross-sectional mean inflation rate deviations of U.S. metropolitan areas and Canadian provinces over the period 1983 to 2002. On the x-axis (denoted by* t*), period's* t *inflation deviations from the cross-regional mean and on the y-axis (denoted by* $t+1$*), period's* $t+1$ *inflation rate 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: Contour Plot of the Estimated Stochastic Kernel and Conditional Expected Next Period's Mean for U.S. and Canadian Regional Mean-Inflation Rate 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 plots expected period's* $t+1$ *mean-inflation rate deviations conditional on period's* t *mean-inflation rate deviations.*

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



Note: *Figure 21 plots kernel density estimates of the empirical distribution of regional inflation rates of our three samples versus the density from a normal distribution that is 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.*

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A Levin-Lin Panel Unit Root Test

A.1 The Test Procedure

To obtain the Levin-Lin panel-unit root results in section 3, we proceed as follows: Let $\pi_{i,t}$ (with $i = 1, 2, \ldots, N$ and $t = 1, 2, \ldots, T$) be a balanced panel of inflation rates consisting of N individual regions with T observations, respectively. The starting point of our analysis is the following test equation:

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

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

$$u_{i,t} = \theta_t + \epsilon_{i,t}. \tag{A.2}$$

In this specification, θ_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 non-stationarity and for the alternative hypothesis of stationarity) the homogeneity restriction that all ρ_i 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 this null hypothesis we proceed as follows:

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

$$\tilde{\pi}_{i,t} = \pi_{i,t} - \frac{1}{N}\sum_{j=1}^{N} \pi_{j,t} \tag{A.3}$$

Having transformed the dependent variable we proceed with the following test equation:

$$\Delta\tilde{\pi}_{i,t} = \rho\tilde{\pi}_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j}\Delta\tilde{\pi}_{i,t-j} + \epsilon_{i,t}. \tag{A.4}$$

The lagged differences of $\tilde{\pi}_{i,t}$ are included to control for potential serial correlations in the idiosyncratic shocks $\epsilon_{i,t}$. Whereas we equalize the ρ_i 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 the general-to-specific method of Hall (1994) which is recommended by Campbell and Perron (1991).

2. The next step in our testing procedure is to run the following two auxiliary

regressions

$$\Delta\tilde{\pi}_{i,t} = \sum_{j=1}^{k_i} \phi_{1i,j}\Delta\tilde{\pi}_{i,t-j} + e_{i,t}. \tag{A.5}$$

$$\tilde{\pi}_{i,t-1} = \sum_{j=1}^{k_i} \phi_{2i,j}\Delta\tilde{\pi}_{i,t-j} + \nu_{i,t-1}. \tag{A.6}$$

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

$$\hat{e}_{i,t} = \rho_i \hat{\nu}_{i,t-1} + \eta_{i,t}. \tag{A.7}$$

The residuals of (A.7) are used to compute an estimate of the variance of $\eta_{i,t}$:

$$\hat{\sigma}^2_{\eta i} = \frac{1}{T - k_i - 1} \sum_{t=k_i+2}^{T} \hat{\eta}^2_{i,t} \tag{A.8}$$

4. Normalizing the OLS residuals $\hat{e}_{i,t}$ and $\hat{\nu}_{i,t-1}$ by dividing them through $\hat{\sigma}_{\eta i}$ yields:

$$\tilde{e}_{i,t} = \frac{\hat{e}_{i,t}}{\hat{\sigma}_{\eta i}} \tag{A.9}$$

$$\tilde{\nu}_{i,t-1} = \frac{\hat{\nu}_{i,t-1}}{\hat{\sigma}_{\eta i}} \tag{A.10}$$

5. The normalized residuals are used to run the following pooled cross-section time-series regression:

$$\tilde{e}_{i,t} = \rho\tilde{\nu}_{i,t-1} + \tilde{\epsilon}_{i,t}. \tag{A.11}$$

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 $\rho = 0$. 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 τ 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 A.2.

A.2 Bootstrap Procedure

Since the finite-sample properties of the adjusted τ statistics are unknown and since idiosyncratic shocks may be correlated across individual regions we rely on bootstrap methods to infer critical values for the τ statistics. More precisely, we employ a nonparametric bootstrap where we resample the estimated residuals from our model. The starting point of our bootstrap approach is given by the hypothesized data generating process (DGP) under the null hypothesis

$$\Delta\pi_{i,t} = \sum_{j=1}^{k_i} \phi_{i,j}\Delta\pi_{i,t-j} + \epsilon_{i,t}. \tag{A.12}$$

Our procedure is as follows:

1. We retrieve the OLS residuals from estimating the DGP under the null hypothesis. This yields the vectors $\hat{\epsilon}_1$, $\hat{\epsilon}_2$, ..., $\hat{\epsilon}_T$, where $\hat{\epsilon}_t$ is the $1xN$ residual vector for period t.

2. Then, we resample these residual vectors by drawing one of the possible T residual vectors with probability $\frac{1}{T}$ for each $t = 1, \ldots, T$.

3. These resampled residual vectors are used to recursively build up pseudo-observations $\Delta\hat{\pi}_{i,t}$ according to the DGP (using the estimated coefficients $\hat{\phi}_{i,j}$).

4. Next, we perform the Levin-Lin test (as described in subsection A.1) on these observations (without subtracting the cross-sectional mean). The resulting τ is saved.

5. Steps two to four are repeated 5,000 times. The collection of the τ statistics form the bootstrap distribution of these statistics under the null hypothesis.

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