

Tillmann, Peter; Wolters, Maik H.

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## The changing dynamics of US inflation persistence: A quantile regression approach

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**No. 06-2012**

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Coordination: Bernd Hayo • Philipps-University Marburg  
Faculty of Business Administration and Economics • Universitätsstraße 24, D-35032 Marburg  
Tel: +49-6421-2823091, Fax: +49-6421-2823088, e-mail: [hayo@wiwi.uni-marburg.de](mailto:hayo@wiwi.uni-marburg.de)

# The changing dynamics of US inflation persistence: a quantile regression approach\*

Peter Tillmann<sup>†</sup>

Justus Liebig University Giessen, Germany

Maik H. Wolters<sup>‡</sup>

Goethe University Frankfurt, Germany

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## Abstract

We examine both the degree and the structural stability of inflation persistence at different quantiles of the conditional inflation distribution. Previous research focused exclusively on persistence at the conditional mean of the inflation rate. Economic theory, however, provides various reasons -for example downward wage rigidities or menu costs- to expect higher inflation persistence at the upper than at the lower tail of the conditional inflation distribution. Based on post-war US data we indeed find slower mean reversion in response to positive than to negative shocks. We find robust evidence for a structural break in persistence at all quantiles of the inflation process in the early 1980s. Inflation persistence has decreased and become more homogeneous across quantiles. Persistence at the conditional mean became more informative about the degree of persistence across the entire conditional inflation distribution. While prior to the 1980s inflation was not mean reverting in response to large positive shocks, our evidence strongly suggests that since the end of the Volcker disinflation the unit root can be rejected at every quantile including the upper tail of the conditional inflation distribution.

Keywords: inflation persistence, quantile regressions, structural breaks, monetary policy, Federal Reserve

JEL classification: E31, E58, C32

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<sup>†</sup>Email: peter.tillmann@wirtschaft.uni-giessen.de

<sup>‡</sup>Email: wolters@wiwi.uni-frankfurt.de (corresponding author)

# 1 Introduction

It is well known that the stance and the strategy of monetary policy in many advanced economies underwent important structural changes over the past decades. Many of these changes can be associated with specific dates in history, such as the early 1980s Volcker disinflation in the US. Other changes appear gradual and are thus more difficult to identify. The changing nature of monetary policy is likely to be reflected in the time series properties of inflation dynamics. While shifts in the mean and the variance of inflation could be interpreted in light of monetary policy changes, another indicator is the degree of persistence, or inertia, in the inflation process. Persistence refers to the speed at which the inflation rate returns to its mean following a shock. Even if the mean inflation rate falls due to the central bank's increased focus on price stability, fluctuations in inflation can be short-lived or long-lasting - in part depending on the determination of monetary policy to bring inflation back on an implicit or explicit target rate. While changes in inflation persistence partly reflect the stance of monetary policy, the degree of persistence contains important information for the conduct of monetary policy, see Gerlach et al. (2009).<sup>1</sup> This is because the higher the degree of inflation persistence, the stronger is the vigour of policy steps necessary to bring inflation back on target.<sup>2</sup>

Hence, the persistence properties of inflation have received considerable attention in the empirical literature. In a recent survey article, Fuhrer (2011, p. 448) summarizes the abundant literature on the nature and the sources of persistent inflation dynamics in the US. He argues

”that the contribution to inflation from its unit root component has diminished significantly in recent decades. [...] With regard to the specific autocorrelation properties of a stationary inflation rate, the picture is considerably murkier.”

We revisit the changing nature of inflation persistence in the US. We add to the literature on inflation persistence in two ways. First, we use a quantile regression approach which allows us to examine the degree of inflation persistence at different conditional quantiles of inflation. Thus far the literature focuses on persistence evaluated at the conditional mean. This neglects the fact that inflation following shocks drawn from the tails of the shock distribution might exhibit a different pattern of inertia than inflation close to the mean. Second, we draw on techniques recently developed by Oka and Qu (2011) to estimate structural changes in regression quantiles

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<sup>1</sup>Bianchi and Ilut (2011) estimate a regime-switching stochastic general equilibrium model for the US economy allowing for a one-time shift from non-Ricardian to a Ricardian regime. They find a shift in the early 1980s that is also able to account for a reduction in inflation persistence.

<sup>2</sup>Other factors such as a reduction in the degree of wage indexation since the early 1980s, see Hoffman, Peersman and Straub (2010), would also lead to a reduction in persistence, although a more gradual decline.

at unknown dates to detect structural changes in persistence for different inflation quantiles. This allows us to examine whether changes in persistence are synchronized across inflation quantiles and whether shifts in persistence at the mean inflation rate are informative about the entire distribution of inflation outcomes.

The asymmetric nature of inflation dynamics depending on the sign of shocks to inflation can be motivated by referring to four different theories: nominal wage rigidity, menu costs, asymmetric monetary policy and regime shifts. We briefly sketch the implications for our analysis.

First, downward wage rigidities have been documented by many researchers. Examples include Akerlof et al. (1996), Card and Hyslop (1998), Dickens et al. (2007), Holden and Wulfsberg (2009) and Messina et al. (2010). Kim and Ruge-Murcia (2009), Abritti and Fahr (2011) and Fahr and Smets (2010) include downward wage rigidities in New Keynesian models through a convex cost function with lower costs for adjusting wages upwards than for cutting them. They find differing transmissions of positive and negative demand shocks from wages to inflation: during expansionary periods real wages and inflation increase considerably, while in contractionary periods shocks are mainly absorbed through a decline in employment with the reaction of inflation being smaller. Sticky prices together with rationing of demand lead to similar asymmetries as downward wage rigidities (see e.g. Cover (1992), Parker and Rothman (2004) and Ravn and Sola (2004)).<sup>3</sup>

Second, menu costs models include -besides asymmetric effects of shocks with different signs- also nonlinearities with respect to the magnitude of shocks.<sup>4</sup> While in the above mentioned literature downward wage rigidities are imposed exogenously, in menu costs models with trend inflation downward rigidities emerge endogenously. Karadi and Reiff (2011) show that large shocks induce more firms to pay menu costs and adjust their prices. Inflation effects become highly nonlinear: while prices are sticky in response to small shocks, prices are endogenously more flexible in response to large shocks due to the adjustment on the extensive margin. Including trend inflation, these differences between small and large shocks lead also to different effects of positive and negative shocks as shown by Caballero and Engle (1993), Tsiddon (1993) and Ball and Mankiw (1994). Firms that want to lower their relative price can simply wait until trend inflation does the work. In contrast, firms will pay the menu cost and adjust prices in response to positive shocks to prevent a gap between the desired relative price and the actual falling relative price. Karadi and Greiff (2011) provide empirical evidence showing that the inflation pass-through is higher for positive than negative shocks.

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<sup>3</sup>Beaudry and Koop (1993) provide reduced-form evidence for the notion that negative output shocks are less persistent than positive shocks. This asymmetry is likely to be translated into different degrees of inflation persistence.

<sup>4</sup>See Ravn and Sola (2004) for reduced-form evidence on the aggregate consequences of nonlinearity in the inflation-effects of shocks that is consistent with the existence of menu costs.

Asymmetric effects of shocks in menu cost models imply that the Phillips Curve is nonlinear. Empirical tests have confirmed such a non-linearity in the inflation-output trade-off implying that positive demand shocks are more inflationary than negative demand shocks (see e.g. Álvarez-Lois (2000), Laxton et al. (1995), Clark et al. (1996), Debelle and Laxton (1997) and Laxton et al. (1999)).

Third, asymmetric monetary policy reactions might also lead to different effects of inflationary shocks depending on the shock size. Policy makers might care more about negative than positive output gaps leading to higher persistence of positive than negative inflationary shocks. Theoretical models include the work by Cukierman (1999), Gerlach (2000) and Nobay and Peel (2003). Gerlach (2000) also provides empirical evidence showing that the Fed exhibited an asymmetric reaction to the output gap prior to 1979, but not afterwards. Ruge-Murcia (2004) finds an asymmetric reaction of the Fed to the output gap for a sample from 1960-1999. Cukierman and Muscatelli (2008) find that a recession avoidance preference was prevailing during the Burns/Miller and the Greenspan periods, while under Volcker the Taylor rule was linear. Ruge-Murcia (1999) shows that the Fed has been more concerned about inflation falling below than rising above the desired level for the period 1952-1999.

Finally, models with switches between different policy regimes could lead to higher inflation persistence in response to large inflation shocks if agents perceive large inflation shocks as an increase in the probability of switching to an inflationary policy regime, while low shocks let them assume that the policy regime will be stable (Bianchi and Melosi, 2011).

In contrast to standard estimates at the conditional mean, our empirical specification allows us to estimate the degree of inflation persistence conditional on different magnitudes and signs of shocks. Given the empirical evidence and the various theoretical models that imply asymmetric effects of inflation shocks of different size, monetary policy makers might be more interested in monitoring the development of inflation persistence at the upper part of the conditional inflation distribution than at the lower part. Standard estimates at the conditional mean cannot distinguish between these important differences in the effects of shocks to inflation. Quantile regressions, in contrast, offer a natural way to empirically assess the importance of these asymmetries. In particular, quantile regressions allow us to analyze whether the timing and the nature of changes of inflation persistence have been synchronous at different parts of its conditional distribution.

Based on a battery of post-war US inflation rates at monthly and quarterly frequency, we derive three key findings:

First, we provide evidence for a structural break in persistence at all quantiles of the inflation process occurring in the early 1950s and, most importantly, in the early 1980s. Persistence at the conditional mean as well as persistence at the outer quantiles is significantly lower after the Volcker disinflation. This result is robust

with respect to changes in mean inflation and the volatility of inflation.

Second, inflation persistence has become more homogeneous across quantiles. Put differently, persistence estimated at the conditional mean of the distribution is highly informative about persistence at other quantiles. Before the 1980s, in contrast, persistence at the outer quantiles was often outside the confidence band surrounding persistence at the mean. This finding also sheds light on previous results in the literature suggesting no change in persistence of deflator inflation. We show that the shift in persistence at the mean rate of deflator inflation is indeed small. At higher quantiles, however, we observe a pronounced drop in persistence.

Third, our evidence strongly suggests that since the end of the Volcker disinflation the unit root can be rejected at every quantile. This sheds light on the recent work on the changing forecastability of inflation. Stock and Watson (2007) model the inflation rate as an integrated moving average, i.e. IMA(1,1), process and argue that the difficulty to forecast inflation rates might stem from changing relative roles of a permanent and a transitory component in the inflation process. In our paper, however, we show that the empirical support for the assumption of a unit root component in the inflation process has disappeared not only at the mean inflation rate but at all quantiles.

The remainder of the paper is organized as follows: Section 2 briefly surveys the available empirical literature. Our empirical approach is presented in section 3. Section 4 introduces the data set and discusses the main results. A battery of robustness checks is documented in section 5. The final section draws some tentative conclusions.

## **2 A brief review of the literature on inflation persistence**

In his survey on inflation persistence, Fuhrer (2011, p. 448) summarizes the empirical evidence: "All authors agree that in the US and many other developed countries inflation exhibited considerable persistence from the 1960s to the mid-1980s. After that time, the statistical evidence is mixed. For both the US and other countries studies fall on both sides of the argument about the possibility of declining reduced-form persistence."

Let us briefly survey the key contributions to the picture described by Fuhrer. In one of the earliest studies, Taylor (2000) finds a break in US inflation persistence that coincides with the Volcker disinflation. Cecchetti and Debelle (2006) and Levin and Piger (2006) assess inflation persistence for major industrial economies and find that conditional on a break in the intercept inflation is much less persistent than previously thought. Both papers stress the need to account for shifts in mean inflation. Neglecting shifts in mean inflation could lead to spuriously high estimates

of the sum of the autoregressive coefficients in the inflation process.<sup>5</sup>

These contributions do not examine structural changes in inflation persistence at potentially unknown points in time. Cogley, Primiceri and Sargent (2010) use a time-varying vector autoregression to estimate the nonstationary trend component of inflation, which they associate with the Fed’s inflation target. Changes in trend inflation could then lead to changes in persistence of aggregate inflation. They find high persistence in inflation inherited from the trend component. This implies that the Fed’s implicit target for inflation continues to have a unit root. They are able to show a reduced persistence in the gap between inflation and the pure random walk component of the inflation rate in the Volcker-Greenspan era.<sup>6</sup> Pivetta and Reis (2007), in contrast, use Bayesian methods and find no change in the persistence of GDP deflator inflation in the US. According to their results based on rolling-window and recursive samples inflation persistence is high and unchanged over the past decades. Benati (2008) systematically evaluates the impact of regime shifts in monetary policy on the persistence properties of inflation. His estimates of the sum of the autoregressive coefficients in a univariate process of CPI inflation drop significantly in the post-Volcker period. Persistence of PCE and deflator inflation, however, remain high even after the Volcker-era.

Evidence on inflation persistence at the level of disaggregate inflation is provided by Clark (2006). His results reveal that the aggregation process induces persistence into the aggregate inflation series despite disaggregate inflation exhibiting little persistence.

An alternative approach to modeling nonlinearities in the persistence properties of inflation is to specify a smooth transition autoregressive (STAR) model where a nonlinear transition function governs the shift between different regimes. Following this line, Nobay, Paya and Peel (2010) find US inflation to be more mean reverting the further away inflation is from its mean. It is important to note that the STAR approach rests on the assumption of a specific functional form of the transition function. Quantile regressions, in contrast, offer a particularly attractive alternative as they do not require a priori assumptions.

A separate strand of the literature examines the degree of fractional integration of the inflation process. Kumar and Okimoto (2007) find a break in the order of fractional integration of monthly US CPI inflation in 1982. Long run persistence—or long memory—of inflation is much lower after the Volcker disinflation period. Recently, Hassler and Meller (2011) extend this line of research and present a test for multiple structural changes in the degree of fractional integration applied to monthly CPI

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<sup>5</sup>In this paper we focus on the US case. O’Reilly and Whelan (2005) show the difficulty to find empirical support for a reduction in persistence in the European Monetary Union.

<sup>6</sup>Kang, Kim and Morley (2009) model regime shifts in persistence in an unobserved components model. They document a shift to the low-persistence regime in the early 1980s. Zhang and Clovis (2009) use formal tests for structural stability of autoregressive models to document shifts in US inflation persistence.



inflation in the US. Conditional on a shift in mean inflation, which they locate in 1981, they find a break in 1973 only. At this date, that coincides with the collapse of the Bretton Woods system and the first oil crisis, inflation persistence significantly increases. A second break in 1980 proves to be insignificant.<sup>7</sup>

Taken together, the literature indeed supports Fuhrer’s (2011) cautious view. The present paper revisits the changing nature of inflation persistence. A potential explanation behind these divergent results could be a sizable degree of heterogeneity of inflation persistence at different conditional quantiles of inflation. If persistence differs according to the size or the sign of the shocks driving inflation, the mean inflation rate would not be informative about the true nature of persistence. To address this issue, we model inflation persistence at different quantiles of the inflation process.

Our study is closely related to the recent work of Tsong and Lee (2011). These authors also model inflation in a quantile framework but do not assess time-variation in persistence at individual quantiles. Our key contribution is to apply recently developed tests for structural breaks at unknown time to the question of inflation persistence.

### 3 A quantile approach to inflation persistence

In this section we introduce the measurement of inflation persistence and sketch the test for structural breaks in persistence at different quantiles.

#### 3.1 Measuring inflation persistence

Our preferred measure of persistence is the sum of the autoregressive coefficients in a univariate process of inflation. By using this reduced form measure of inflation persistence we do not take a stand on the structural sources of persistence. A change in persistence detected by our measure is consistent with a variety of structural changes in the conduct of monetary policy, the nature of nominal rigidities or the properties of shocks hitting the economy. Let  $\pi_t$  be the inflation measure,  $\alpha$  an intercept term and  $\varepsilon_t$  a serially uncorrelated error term. The AR( $q$ ) process is

$$\pi_t = \alpha + \sum_{k=1}^q \beta_k \pi_{t-k} + \varepsilon_t \quad (1)$$

The sum of autoregressive coefficients is  $\rho = \sum_{k=1}^q \beta_k$ . According to Andrews and Chen (1994),  $\rho$  is the preferred scalar measure of persistence in  $\pi_t$ , since a monotonic

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<sup>7</sup>Since the long memory property of inflation examined in these studies can be approximated by an autoregressive process of very high order, a change in the degree of fractional integration can be interpreted as a shift in inflation persistence over the very long run. The approach taken in this paper, however, focuses on autoregressive processes of lower order.

relationship exists between  $\rho$  and the cumulative impulse response function of  $\pi_{t+j}$  to  $\varepsilon_t$ . Rewrite expression (1) as

$$\pi_t = \alpha + \rho\pi_{t-1} + \sum_{k=1}^{q-1} \gamma_k \Delta\pi_{t-k} + \varepsilon_t \quad (2)$$

where  $\Delta\pi_t = \pi_t - \pi_{t-1}$ . If  $\rho = 1$ , the inflation process contains a unit root. If  $|\rho| < 1$ , the process is stationary. In the empirical application below we set the lag length to  $q = 4$  for quarterly data and  $q = 6$  for monthly data.

Estimates of  $\rho$  obtained from least squares estimation suffer from a bias as  $\rho$  approaches unity. Therefore, the literature typically resorts to Hansen's (1999) median unbiased estimator of  $\rho$ . This estimator, however, has not yet been developed for quantile autoregressive models. We follow Tsong and Lee (2011) and, when referring to estimates for conditional quantiles, report results based on the standard quantile regression estimates by Koenker and Bassett (1978).

### 3.2 Persistence at different quantiles

Quantiles are values that divide a distribution such that a given proportion of observations is located below the quantile. The  $\tau - th$  quantile is defined as the value  $q_\tau(\pi_t|\pi_{t-1}, \dots, \pi_{t-q})$  such that the probability that the conditional inflation rate will be less than  $q_\tau(\pi_t|\pi_{t-1}, \dots, \pi_{t-q})$  is  $\tau$  and the probability that it will be more than  $q_\tau(\pi_t|\pi_{t-1}, \dots, \pi_{t-q})$  is  $1 - \tau$ . The AR( $q$ ) process of inflation dynamics at quantile  $\tau$  can be written as a quantile autoregression, QAR( $q$ )

$$q_\tau(\pi_t|\pi_{t-1}, \dots, \pi_{t-q}) = \alpha(\tau) + \rho(\tau)\pi_{t-1} + \sum_{k=1}^{q-1} \gamma_k(\tau) \Delta\pi_{t-k} \quad (3)$$

Estimating the persistence parameter at different quantiles of the distribution instead of the mean can be done with quantile regressions as introduced by Koenker and Bassett (1978). Following the work of Koenker and Xiao (2004, 2006), this gives us the persistence parameter conditional on a grid of values for  $\tau$ . Quantile regressions impose no functional form constraints on parameter values over the conditional distribution of the inflation rate.

The interpretation of the quantile regression approach to inflation persistence is straightforward: estimates of  $\rho(\tau)$  reveal the extent of inflation persistence at the quantile  $\tau$  conditional on past values of inflation  $\pi_{t-1}, \dots, \pi_{t-q}$ . Thus, shocks to the inflation process of different size and magnitude are allowed to lead to different patterns of persistence. Strictly speaking, however, we cannot interpret persistence at, say, the  $\tau = 0.1$  quantile as reflecting persistence at low absolute levels of inflation. Rather, it measures persistence when inflation exhibits a large negative deviation from its mean.

### 3.3 Breaks in persistence at different quantiles

We test for structural breaks in inflation persistence that might show up in any part of the conditional inflation distribution. Qu (2008) and Oka and Qu (2011) have developed tests for structural change with unknown timing in regression quantiles. The tests are subgradient based and have good properties in small samples.

The test is run in two stages as recommended by Qu (2008) and Oka and Qu (2011). First, we test for structural stability across a range of quantiles using the  $DQ$ -test. This is a general test for changes in the entire conditional distribution of inflation. Since we do not have any prior information as to which part of the conditional inflation rate distribution is subject to a break we test for a large range of  $\tau \in \{0.05, 0.1, \dots, 0.95\}$ . The disadvantage of using a wide range is that the power of the test decreases as opposed to the case where prior information is used to trim the range of quantiles. Therefore, in a second step we test for structural change in prespecified quantiles using the  $SQ_\tau$ -test. If the  $DQ$ -test rejects the null hypothesis of no structural break, the  $SQ_\tau$ -test can reveal structural breaks in different parts of the conditional inflation rate distribution. In this way we can detect in which parts of the distribution the actual change takes place and obtain a full picture about the stability of persistence across quantiles.

The tests allow for multiple structural breaks with unknown timing. The test procedure runs sequentially: first, for a given number of breaks, the break dates and the AR parameters are estimated jointly by minimizing the quantile check function over all permissible break dates. The range of permissible break dates excludes the first and the final 5% of the observations. We repeat this procedure for one to a maximum of ten possible structural breaks. Second, we use the  $DQ$ - and  $SQ_\tau$ -test to consider the existence of one structural break against the null hypothesis of no structural break.

Let  $\xi(\tau) = (\alpha(\tau), \rho(\tau), \gamma_1(\tau), \dots, \gamma_{q-1}(\tau))$  denote the vector of parameters in equation (3) at quantile  $\tau$  and suppose our inflation series contains  $T$  observations. Then the hypotheses for the  $DQ$ -test are:

$$\begin{aligned} H_0^* : \xi_i(\tau) &= \xi_0(\tau) \text{ for all } i \text{ and for all } \tau \in \{0.05, 0.1, \dots, 0.95\} \\ H_1^* : \xi_i(\tau) &= \begin{cases} \xi_1(\tau) & \text{for } i = 1, 2, \dots, t \\ \xi_2(\tau) & \text{for } i = t + 1, \dots, T. \end{cases} \text{ for some } \tau \in (0, 1) \end{aligned}$$

The two hypotheses for the  $SQ_\tau$ -test are given by:

$$\begin{aligned} H_0 : \xi_i(\tau) &= \xi_0(\tau) \text{ for all } i \text{ for a given } \tau \in \{0.05, 0.1, \dots, 0.95\} \\ H_1 : \xi_i(\tau) &= \begin{cases} \xi_1(\tau) & \text{for } i = 1, 2, \dots, t \\ \xi_2(\tau) & \text{for } i = t + 1, \dots, T. \end{cases} \end{aligned}$$

First, we estimate the AR parameters at the different conditional quantiles under the null of no structural break. Afterwards, we estimate the AR parameters at the different conditional quantiles separately for the subsamples based on the previously estimated break date. If a structural break exists, the estimated parameters under the null hypothesis are not close to the true values for at least one subset of the sample. The estimated residuals will persistently fall below (or above) the true quantile, forcing the subgradient to take a large value.

If the null hypothesis of no structural change is rejected, we test in sequential steps the null hypothesis of 1 break against the alternative hypothesis of 2 breaks using the  $DQ(l+1|l)$ - and the  $SQ_\tau(l+1|l)$ -tests. If we find evidence in favor of 2 breaks we check the null hypothesis of 2 breaks against the alternative hypothesis of 3 breaks and so on. Tables for critical values are provided in Qu (2008).

## 4 Data and results

### 4.1 The data set

We use measures of post-war US inflation based on three alternative price indices with two different frequencies. The first measure is inflation expressed as the annualized quarter-on-quarter or month-on-month percentage change of the Consumer Price Index (CPI) for all urban consumers. The second measure is the annualized quarter-on-quarter or month-on-month percentage change of the Personal Consumption Expenditure (PCE) chain-type price index. Our third measure is the quarter-on-quarter percentage change of the GDP deflator. All data series are taken from the FRED database at the Federal Reserve Bank of St. Louis. The resulting five inflation series are depicted in figure (1). Table (1) presents some descriptive statistics for each series and table (2) reports the correlation coefficients among the quarterly and monthly series, respectively. Interestingly, the correlation between CPI inflation on the one hand and PCE or deflator inflation on the other weakened substantially in the post-1984 period.<sup>8</sup> This also suggests that changes in persistence, if any, could be unevenly spread across alternative inflation rates.

### 4.2 Results

As a first step, we study the behavior of inflation persistence at the conditional median (solid line) and mean (dotted line, without confidence band) of the series. Figures (2) and (3) present rolling-window estimates of persistence together with bootstrapped confidence bands for a 10-year window. The results reflect the consensus view portrayed before: inflation has become less persistent since the early 1980s. This tendency is more pronounced for monthly inflation rates and, in par-

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<sup>8</sup>See Hakkio (2008) for a deeper analysis of the differences between PCE and CPI inflation in terms of weights, computation and scope.

ticular, for CPI inflation. Inflation persistence based on the GDP deflator, instead, exhibits fewer signs of instability, which is consistent with Pivetta and Reis’s (2007) finding and Benati’s (2008) result that persistence of CPI inflation has decreased, while it has remained high for PCE and deflator inflation. This reflects the change in the correlation structure among alternative inflation indicators discussed before. Prior to the early 1980s, when the correlation between all three measures of inflation was high, persistence behaved similarly. Since then, however, persistence diverged across alternative series.

Moreover, some confidence bands still contain the unit root case. In order to assess whether this finding is informative for the entire range of inflation quantiles, we now turn to the results from quantile regressions.<sup>9</sup> To test for the existence of structural breaks in persistence at different quantiles, we report results for the DQ-test introduced in the previous section. Table (3) shows the estimated break dates over the conditional inflation distribution ( $\tau \in \{0.05, 0.1, \dots, 0.95\}$ ). Two findings stand out. First, the break points differ across the alternative inflation rates. For deflator inflation and monthly PCE inflation we do not find structural breaks in the conditional distribution. Quarterly PCE inflation exhibits only a very recent break in 2008Q3. The most persuasive evidence is available for breaks in the distribution of CPI inflation. The differences across inflation measures do not come as a surprise. Since for most of the sample period the Federal Reserve uses CPI inflation as the primary indicator of price pressure, shifts in monetary policy will be most directly reflected in CPI inflation. Second, besides the breaks in the early 1950s, the Volcker disinflation in 1981 is clearly associated with a break in the entire distribution of quarterly and monthly CPI inflation.

Because the previous test for a joint break at all quantiles might be too restrictive, figures (4) and (5) report the results of the  $SQ_\tau$ -test for breaks at specific quantiles for the quarterly and monthly inflation series, respectively. Structural breaks in specific quantiles occur frequently. However, PCE inflation exhibits breaks only at the upper quantiles while deflator inflation is subject to breaks only at very low quantiles. Only CPI inflation exhibits breaks over the entire range of quantiles (see figure (4)). While there is also evidence for a break in CPI inflation in the early 1950s, the breaks occurring in the 1980s are more interesting for our purposes. The results for the monthly inflation rates presented in figure (5) provide strong evidence for structural breaks in CPI inflation at all quantiles in the early 1980s. Breaks in PCE inflation are scattered through the entire sample and concentrated on the outer quantiles only. The shift in monetary policy under Paul Volcker had

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<sup>9</sup>The figures show that quantile regressions at the conditional median, i.e. least absolute deviation estimates, are useful complements to least squares regressions even if one is only interested in characterizing the location of the conditional distribution rather than estimates in the tails of the conditional distribution: the sharp deflationary spike in 2008Q4 leads to a decrease in estimated inflation persistence at the conditional mean (Hansen’s (1999) unbiased median estimator), but not at the conditional median. The latter estimate is robust against outliers.

a profound impact on the inflation process leading to a break in persistence at all conditional quantiles. The weaker evidence for a regime shifts in inflation dynamics for other inflation series reflects a general difficulty. Based on time series evidence a structural break is often difficult to locate despite convincing narrative evidence for the existence of a policy change. This points to the gradual effect of policy changes on observed inflation dynamics.<sup>10</sup>

While we see clear evidence for structural breaks, we do not know both the size and the sign of these shifts yet. To illustrate the nature of these breaks, figures (7) to (11) plot the estimated constant  $\alpha(\tau)$  and the persistence parameter  $\rho(\tau)$  together with bootstrapped 95% confidence bands at different quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  for different subsamples. The subsamples are chosen based on the break points as suggested by the DQ-test for CPI inflation. The first row of each figure presents the quantile-specific estimate of the constant across subsamples, the second row depicts the estimated sum of the autoregressive coefficients.

In all figures the level-shifts in persistence across subsamples is apparent. After the break in 1951 inflation persistence evaluated at the mean inflation rate shifts upwards. Following the break in 1981, in contrast, persistence unanimously falls. This fall is more pronounced for CPI inflation and less clear for PCE and deflator inflation.<sup>11</sup> Our quantile regressions reveal that the fall in persistence is indeed a characteristic of all inflation quantiles. Not only mean inflation became significantly less inertial, but also inflation following particularly large shocks, either negative or positive ones. Moreover, since the break in 1981, persistence is more homogeneous across quantiles. A larger degree of homogeneity across quantiles translates into a smaller degree of nonlinearity with respect to the effects of shocks of different size. Consider the estimated persistence before 1981. For all inflation series persistence at the outer quantiles, e.g.  $\tau = 0.1$  and  $\tau = 0.9$ , significantly deviates from persistence at the conditional mean. After 1981, however, persistence across the entire distribution lies within the confidence band around the persistence estimate obtained at the conditional mean. Thus, the results in the literature, which are exclusively derived from persistence at the mean of the conditional distribution, are informative about the entire inflation distribution.

We find that persistence is generally larger for upper quantiles of the conditional distribution of inflation. Hence, large shocks to inflation generate stronger inertia than smaller shocks. This asymmetric nature of inflation persistence across conditional quantiles is difficult to reconcile with standard model of time-dependent pricing along the lines of Calvo (1983). Menu cost models as pioneered by Ball and Mankiw (1994), in contrast, provide some theoretical underpinning to the observation that the degree of inflation persistence depends on the size and the sign of the

<sup>10</sup>Wieland and Wolters (2011), for example, show that forecasters overestimated inflation following the Volcker disinflation until the 1990s.

<sup>11</sup>This is consistent with evidence provided by Zhang and Clovis (2009).

shocks to the inflation process. Karadi and Reiff (2011) analyze price setting behavior in a menu cost model with fat tailed idiosyncratic shocks. They show that due to additional at the extensive margin, i.e. the number of firms adjusting prices, the inflation response to shocks is nonlinear in the shock size and asymmetric between positive and negative shocks.

As mentioned before, Pivetta and Reis (2007) and Benati (2008) do not find a shift in persistence of deflator inflation. Our methodology sheds light on this result. Figure (9) shows that the fall in persistence at the mean is indeed small. Evaluated at higher quantiles, however, the drop in persistence is much stronger. For  $\tau = 0.9$ , for example, persistence prior to 1981 is significantly above the already high persistence estimate at the mean. After 1981, in contrast, persistence at  $\tau = 0.9$  no longer deviates from persistence at the mean.

Finally, the results shed light on the unit root behavior of inflation. For the most recent subsample we can exclude the unit root in the inflation process at all conditional quantiles. The  $\rho = 1$  case is no longer covered by the bootstrapped confidence bands. Even in the aftermath of large shocks inflation is mean reverting.<sup>12</sup>

## 5 Robustness

### 5.1 The role of the lag order

We repeat the structural break tests for monthly data with a specification with 12 lags. With 12 monthly lags this specification covers dynamics of a whole year as in the case of quarterly data where we have used four lags. The DQ-test detects in this case 3 breaks on the 5% significance level: 1952M9, 1981M9 and 2006M9. In the baseline specification we found two breaks in the 1950s (1951M2 and 1954M7), no break in 2006, but exactly the same break in 1981M9. This additional specification shows that the breakpoint in 1981 is a robust finding, while uncertainty is higher for other potential breakpoints. For a specification with PCE inflation and 12 monthly lags we do not find any break on the 5% level as in the baseline case. Figure 6 shows the results of the SQ-test for monthly CPI and monthly PCE based on the specification with 12 lags. The results are similar to the baseline estimates. The break in 1981 occurs at all quantiles, while breaks in the 1950s and around 2000 are only visible for specific parts of the conditional inflation distribution.

### 5.2 The role of changes in mean inflation

The presence of structural breaks in mean inflation affects the estimates of persistence. Neglecting a break could lead to spuriously high estimates of the sum of the

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<sup>12</sup>Tsong and Lee (2011) also find a unit root for larger inflation quantiles. Our results suggest that this finding disappears once we allow for structural breaks.

autoregressive coefficients. In their analysis of changes in the degree of fractional integration of US inflation, Hassler and Meller (2011) demean the inflation rate before and after 1981 separately to control for a level shift. To corroborate the robustness of our results, we follow Hassler and Meller (2011) and subtract the mean from the CPI inflation series separately for each subsample as identified by the DQ-test. We then employ again the DQ-test to detect breaks in the entire conditional distribution. The breaks found, see table (4), are exactly the same as in the baseline case except that for monthly CPI an additional break in 2000M11 is found. The changes in parameters based on this additional break are modest. These findings support the notion that the baseline results are not obscured by structural breaks in the mean inflation rate. The results for individual quantiles, see figures (12) to (14), are unchanged. The inflation process changed at all conditional quantiles in 1951 and 1981. Furthermore, all other results remain valid. Note also that the upward shift in the intercept term in 1981 that plagued the estimates for CPI inflation presented before disappeared. Taken together, accounting for shifts in the mean inflation rate does not affect our main results. In addition we check the monthly CPI specification with 12 lags. After controlling for mean shifts we find exactly the same three breaks as before: 1952M9, 1981M9 and 2006M9.

### 5.3 The role of changes in inflation volatility

In addition to mean shifts, changes in the volatility of inflation might have an effect on the results. Inflation volatility dropped considerably since the early 1980s. Thus, there is the possibility that the tests by Oka and Qu (2011) falsely detect a change in inflation persistence if in fact only a change in inflation volatility has occurred. To assess this possibility we simulate time series with constant persistence, i.e. constant AR parameters, but a change in the variance of shocks. To generate those series, we estimate AR parameters of quarterly CPI inflation at the conditional mean over the whole sample from 1947 to 2011. Based on these parameters we construct artificial series of the same sample length as quarterly CPI inflation. We calibrate the shock variance separately for the identified subsamples from table (3) to match the inflation volatility in these subsamples. In this way we create 1000 quarterly series with constant persistence, but structural breaks in the variance occurring at the observation numbers that correspond to 1951Q3 and 1981Q3.<sup>13</sup> The variance of annualized quarterly CPI inflation is 33.28 for the first subsample, 14.18 for the second subsample and 4.22 for the third subsample. We run the DQ-test to check whether it falsely detects structural breaks in persistence. On the 5% significance level no structural break is found in 61%, 1 structural break is found in 35% and two breaks are found in 4% of the structural break test simulations. Of the breaks, 41% occur in a four year window around 1951Q3, 16% occur in a four year window

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<sup>13</sup>The series that we generate are 100 observations longer than the sample of quarterly CPI. We eliminate the first 100 observations to make sure that initial conditions do not influence the results.



around 1981Q3, 42% occur in between these dates and 1% after 1983Q3. These results show that in the majority of cases (in 61% in our simulation exercise) the test will not falsely detect a structural break in persistence if in fact a break in volatility has occurred, but the possibility cannot completely be excluded especially for the large inflation volatility drop in the early 1950s that led to a number of falsely detected breaks in the simulated time series. The danger of falsely detecting a break in the early 1980s is much lower than falsely detecting a break in the early 1950s.

These results are confirmed by simulating series that correspond to monthly CPI inflation data. We simulate 100 series<sup>14</sup> of the same sample length as monthly CPI with constant AR coefficients, but with breaks in inflation volatility at the observations that correspond to 1951M2, 1954M7 and 1981M9 (see table 3). We calibrate the shock variance so that the simulated series match the variance of annualized monthly CPI inflation for the four subsamples: 70.59, 9.78, 17.93 and 10.14. In 66% there is no structural break detected, in 31% one structural break is detected and in 3% of the simulations 2 structural breaks are detected. Of these 37 detected breaks 34 occur at observations that correspond to dates between 1950 and 1952 and only one break is detected at the observations that correspond to 1979. The two remaining breaks are at observations that correspond to 1988 and 1991. Thus, based on monthly data the risk of falsely detecting a structural break in persistence if in fact this break is caused by a change in inflation volatility is present for the breaks in the 1950s. The risk of falsely detecting a break in 1980, however, is extremely low.

#### 5.4 The role of changes in the mean and the volatility

In an additional step, we go beyond taking into account only mean shifts and normalize quarterly and monthly CPI inflation in the subsamples identified before (table 3) by their respective standard deviations. We thus normalize the data separately for each subsample to eliminate changes in the mean and the variance. We apply the DQ-test and the  $SQ_\tau$ -tests to these normalized inflation series and compute parameter estimates over the whole conditional inflation distribution for the different baseline subsamples. The DQ-test has difficulties in detecting the breaks that we found previously. On the 5% significance level it only finds structural breaks in inflation persistence in 2000Q3 and 2000M12, for quarterly and monthly data respectively. The  $SQ_\tau$ -test results presented in figure (15) show, however, that in fact over a large range of the conditional inflation distribution structural breaks occur in the early 1980s even after normalizing the data separately for each subsample. There is, however, no evidence for structural breaks in the early 1950s. We repeat the DQ-test for the 10% significance level. For quarterly CPI inflation again only a break in 2000Q3 is detected. For monthly CPI inflation in addition to the break in

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<sup>14</sup>The simulations are extremely time-consuming such that it is hardly feasible to run 1000 monthly simulations.

2000M11 also the previously found break in 1981 is detected (the test shows a break in 1981M11). The parameter estimates at individual quantiles for the different subsamples from table (3), see figures (16) to (17), still show a clear decline in inflation persistence in the early 1980s and also evidence for changes in inflation persistence in the 1950s. Repeating the DQ-test for monthly CPI data with 12 lags, we find evidence for a structural break in the early 1980s on the 5% significance level even after controlling for shifts in the mean and in volatility. The breaks detected in this case are: 1982M4, 1987M1 and 2000M1.

While the simultaneous change in inflation persistence and inflation volatility complicates the analysis, our careful examination shows clear evidence for a structural break in CPI inflation persistence, even though the DQ-test has difficulties for some specifications to detect it when normalizing the data based on subsamples. Based on the results from the  $SQ_\tau$ -test, the visual assessment of parameter estimates for the different subsamples and the DQ-test for monthly CPI inflation with 12 lags, we can exclude the possibility that our results falsely detect changes in inflation persistence that are in fact caused by mean and variance shifts.

## 6 Concluding remarks

We draw on recently developed methods to identify structural breaks at conditional quantiles to study the changing nature of US inflation persistence. The framework is flexible enough to allow for asymmetries of inflation persistence across inflation quantiles - a characteristic of the data that accords well to several theoretical foundations.

We find strong and robust evidence for a reduction in persistence at all conditional quantiles of inflation with persistence becoming more homogeneous across quantiles. Thus, we contributed to the literature on inflation persistence by providing a missing key element: when there are shifts in monetary policy not only persistence at the conditional mean changes. Rather, the results support the notion that the entire inflation process reflects shifts in monetary policy. While Benati (2008) documents that shifts in monetary policy reduce persistence, we show that the new monetary policy regime in the US left its footprint on the entire distribution of inflation.

Our results add to our knowledge about the conduct of monetary policy in at least two dimensions. First, the reduction in persistence is consistent with monetary policy successfully stabilizing inflation around the mean. Even shocks drawn from the tails of the distribution have only a short-lived impact on inflation as monetary policy keeps the inflation rate under control. Second, we question the unit root property of inflation that is often used to decompose inflation into a permanent and a transitory component. It remains to be seen how the recent shift towards unconventional monetary policy measures since 2008 or the prolonged period of

very low inflation rates are reflected in the conditional distribution of inflation. We leave that issue for future research.

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Table 1: Summary statistics

Series	1960-2010	1960-1984	1985-2010	1995-2010
	Means			
CPI (q)	3.98	5.16	2.84	2.40
PCE (q)	3.55	4.66	2.48	2.01
GDP (q)	3.55	4.79	2.37	2.04
CPI (m)	3.96	5.13	2.84	2.40
PCE (m)	3.54	4.64	2.48	2.00
	Standard Deviations			
CPI (q)	3.14	3.63	2.05	2.19
PCE (q)	2.60	2.99	1.54	1.53
GDP (q)	2.41	2.81	1.01	0.94
CPI (m)	3.86	4.15	3.19	3.55
PCE (m)	3.04	3.27	2.36	2.53

Notes: Means and standard deviations for different quarterly (q) and monthly (m) inflation series and alternative subsamples.

Table 2: Correlation matrix

Quarterly series:							
1960-1984				1985-2011			
	CPI	PCE	GDP		CPI	PCE	GDP
CPI	1.00			CPI	1.00		
PCE	0.95	1.00		PCE	0.95	1.00	
GDP	0.90	0.94	1.00	GPD	0.58	0.69	1.00
Monthly series:							
1960-1984				1985-2011			
	CPI	PCE			CPI	PCE	
CPI	1.00			CPI	1.00		
PCE	0.89	1.00		PCE	0.40	1.00	

Notes: Correlation coefficients of different quarterly and monthly inflation series for two subsamples.

Table 3: Tests for structural breaks in regression quantiles (DQ-Test)

Series	1st break date	2nd break date	3rd break date	4th break date
CPI (q)	1951Q3	1981Q3	-	-
PCE (q)	2008Q3	-	-	-
GDP defl (q)	-	-	-	-
CPI (m)	1951M2	1954M7	1981M9	-
PCE (m)	-	-	-	-

Notes: Results for the DQ-Test for breakpoints over the conditional inflation distribution ( $\tau \in \{0.05, 0.1, \dots, 0.95\}$ ) on the 5% significance level.



Table 4: Tests for structural breaks in regression quantiles with demeaned data (DQ-test)

Series	1st break date	2nd break date	3rd break date	4th Break date
CPI (q)	1951Q3	1981Q3	-	-
CPI (m)	1951M2	1954M7	1981M9	2000M11

Notes: The table shows results for the DQ-Test for breakpoints over the conditional inflation distribution ( $\tau \in \{0.05, 0.1, \dots, 0.95\}$ ) on the 5% significance level after controlling for a mean shift.

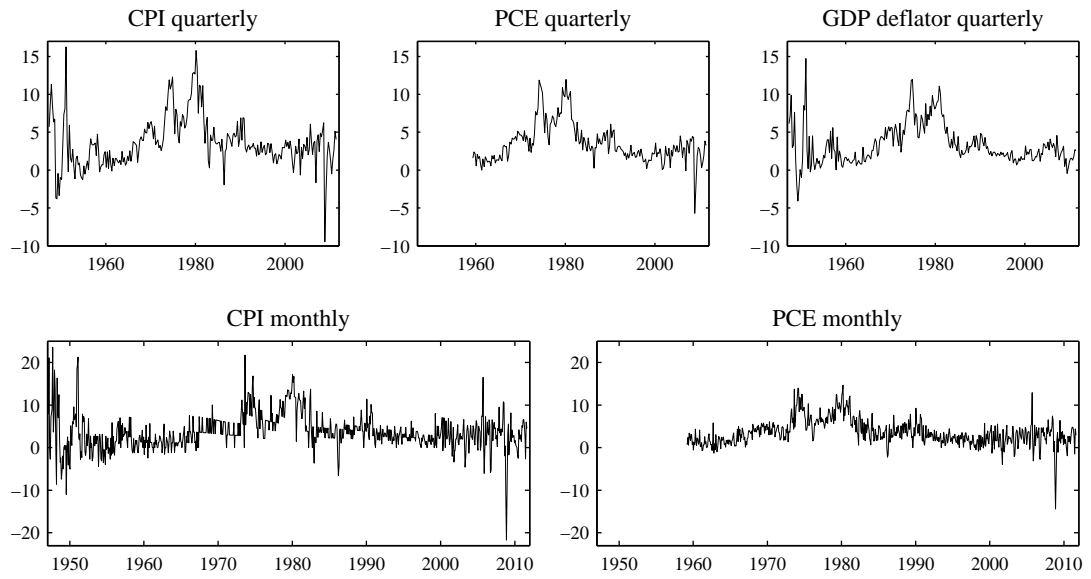


Figure 1: Inflation series. Notes: the different plots show annualized quarter-on-quarter and month-on-month inflation series. CPI: based on the Consumer Price Index for all urban consumers (all items), PCE: based on the Personal Consumption Expenditures chain-type price index, GDP deflator: based on the Implicit Price Deflator of GDP.

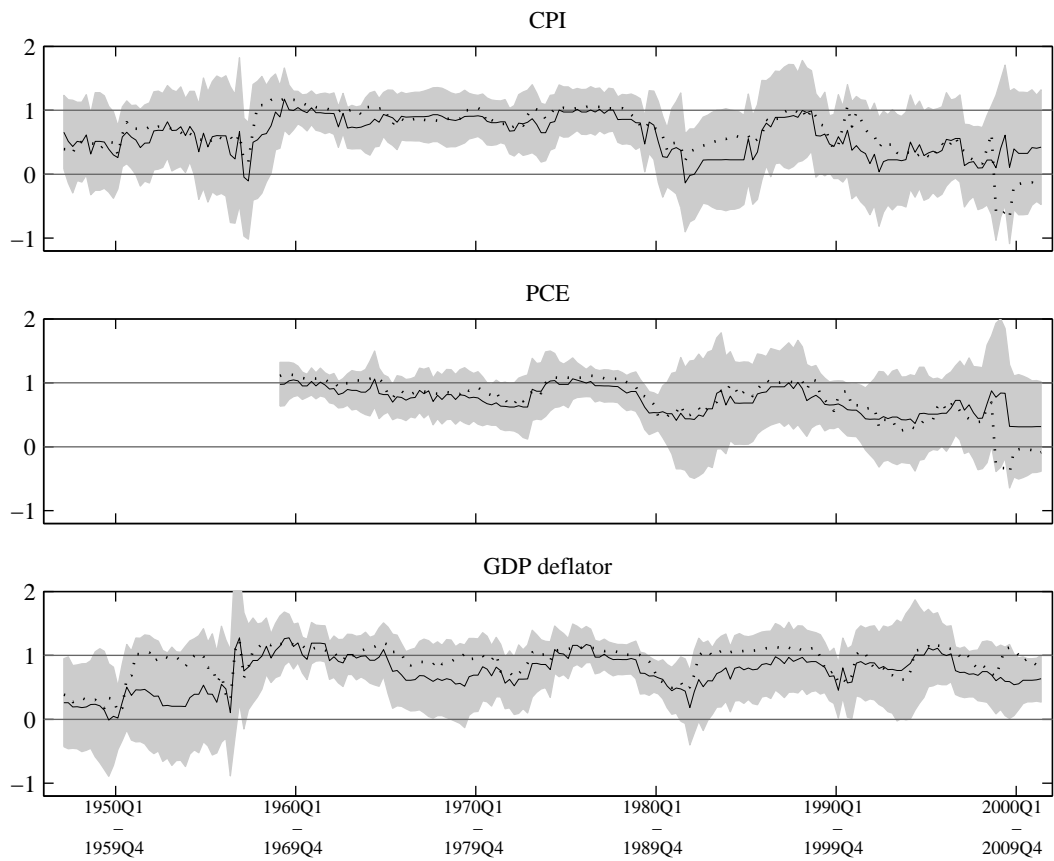


Figure 2: Inflation persistence (quarterly). Notes: The graphs show 10-year rolling window estimates of  $\rho$  (see equation (2)) at the conditional median (solid line) together with 95% bootstrapped confidence bands (gray areas). The dotted line shows point estimated computed with Hansen's (1999) median unbiased estimator.

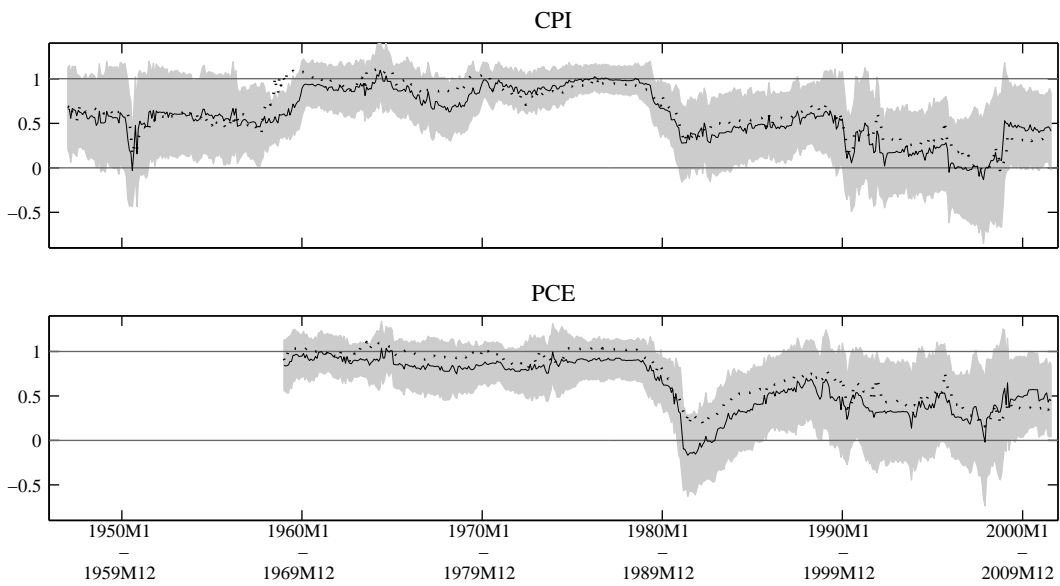


Figure 3: Inflation persistence (monthly). Notes: see figure 2.

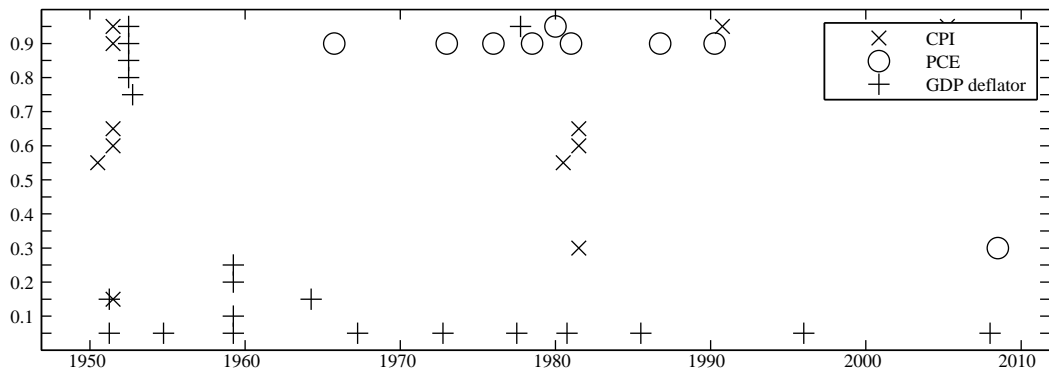


Figure 4: Estimated break points at different quantiles (quarterly). Notes: The graph shows breakpoints estimated with the  $SQ_\tau$ -test at all quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  (vertical axis) for CPI, PCE and GDP deflator inflation.

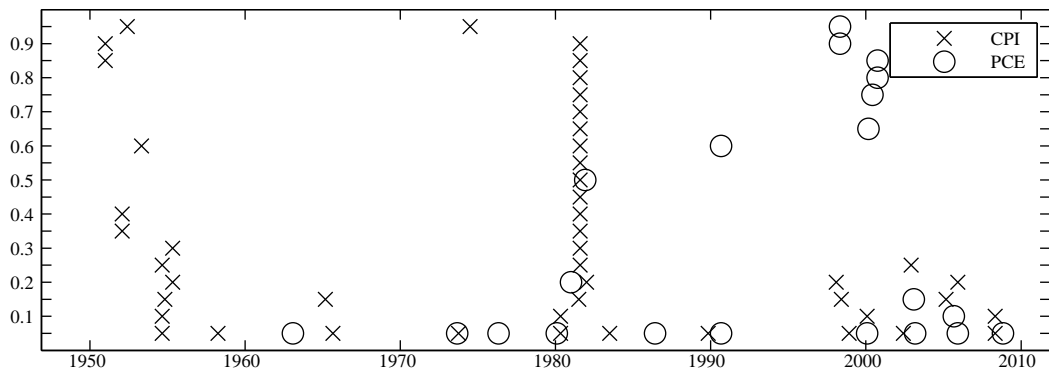


Figure 5: Estimated break points at different quantiles (monthly). Notes: The graph shows breakpoints estimated with the  $SQ_\tau$ -test at all quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  (vertical axis) for CPI and PCE inflation.

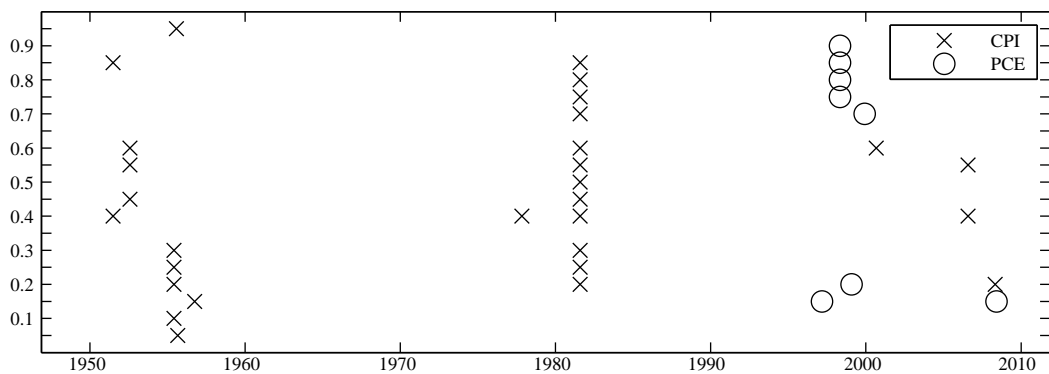


Figure 6: Estimated break points at different quantiles (monthly, 12 lags). Notes: The graph shows breakpoints estimated with the  $SQ_\tau$ -test at all quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  (vertical axis) for CPI and PCE inflation with a specification with 12 lags.

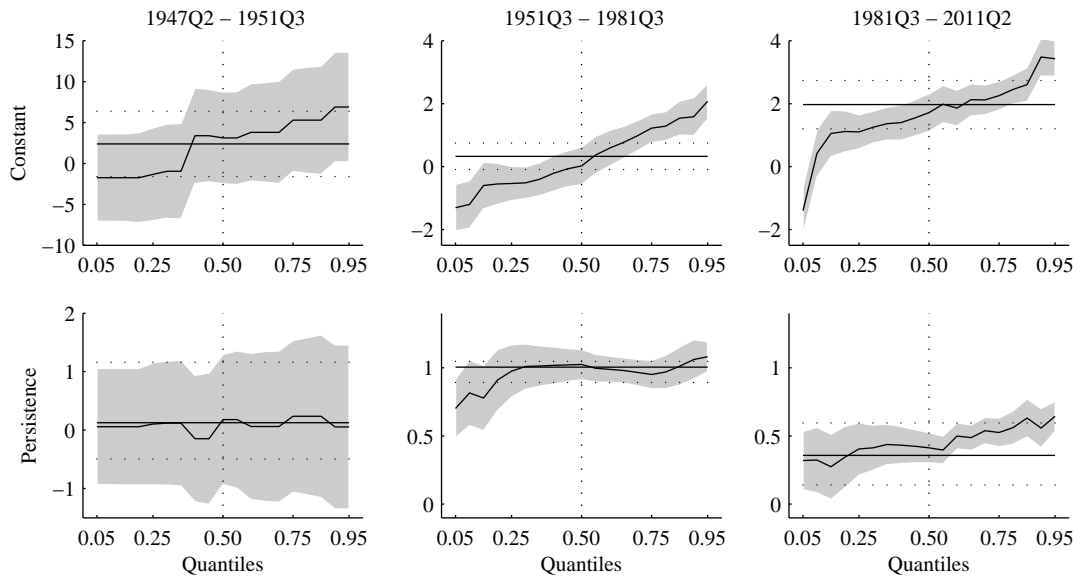


Figure 7: Estimated parameters at different quantiles for quarterly CPI inflation. Notes: The graphs show estimates of the constant  $\alpha$  and the persistence parameter  $\rho$  at different quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  for different subsamples. The gray areas indicate 95% bootstrapped confidence bands. The subsamples are chosen based on the estimated break points of the DQ-test for CPI inflation. The horizontal line shows Hansen's (1999) median unbiased estimator together with 95% grid-t bootstrap confidence bands.

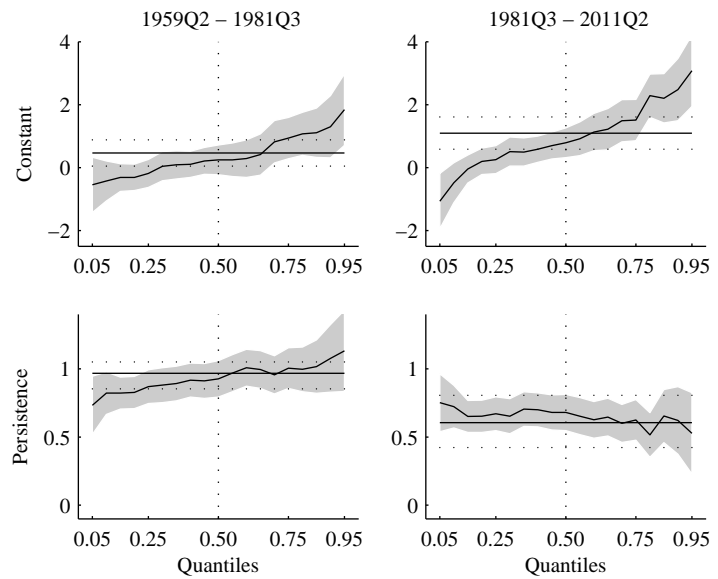


Figure 8: Estimated parameters at different quantiles for quarterly PCE inflation. Notes: The graphs show estimates of the constant  $\alpha$  and the persistence parameter  $\rho$  at different quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  for different subsamples. The gray areas indicate 95% bootstrapped confidence bands. The subsamples are chosen based on the estimated break points of the DQ-test for CPI inflation. For PCE inflation we can only plot estimates for the last two regimes as the data series starts later.

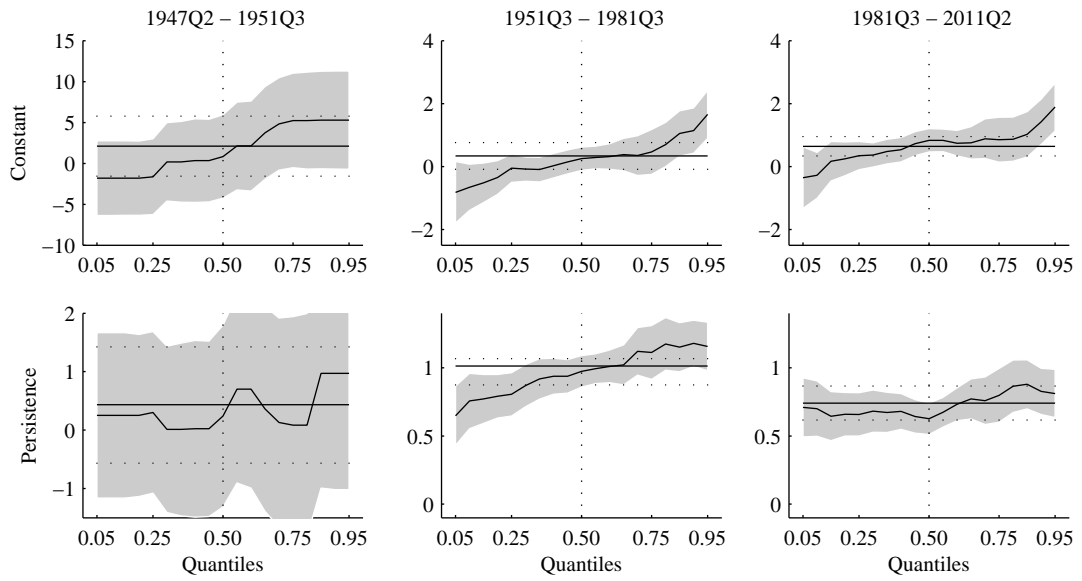


Figure 9: Estimated parameters at different quantiles for quarterly GDP deflator inflation. Notes: The graphs show estimates of the constant  $\alpha$  and the persistence parameter  $\rho$  at different quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  for different subsamples. The gray areas indicate 95% bootstrapped confidence bands. The subsamples are chosen based on the estimated break points of the DQ-test for CPI inflation.

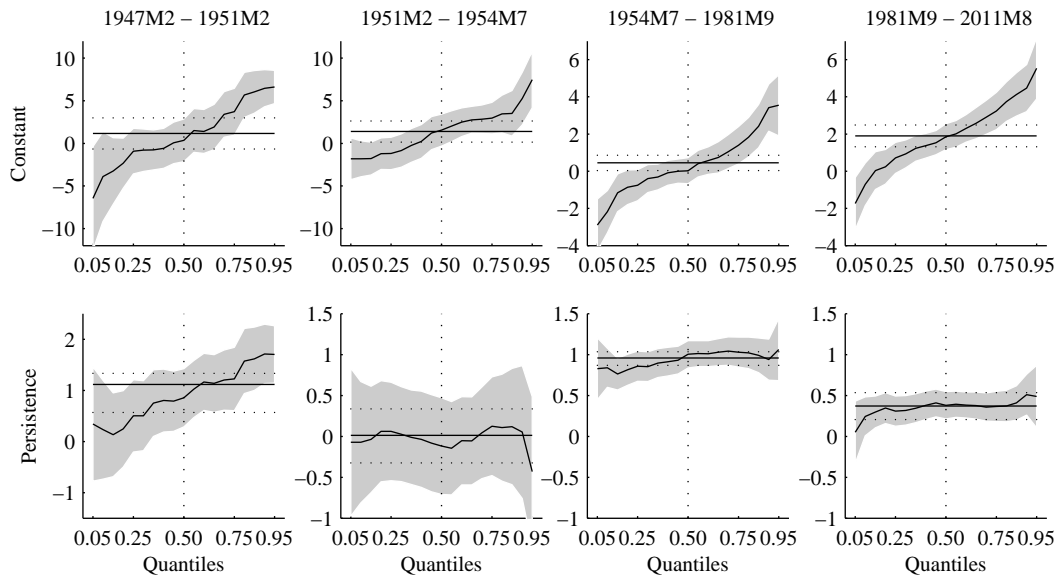


Figure 10: Estimated parameters at different quantiles for monthly CPI inflation. Notes: The graphs show estimates of the constant  $\alpha$  and the persistence parameter  $\rho$  at different quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  for different subsamples. The gray areas indicate 95% bootstrapped confidence bands. The subsamples are chosen based on the estimated break points of the DQ-test for CPI inflation.

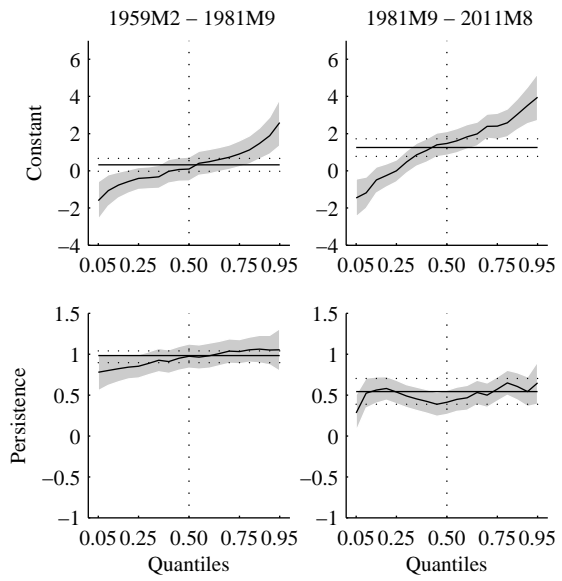


Figure 11: Estimated parameters at different quantiles for monthly PCE inflation. Notes: The graphs show estimates of the constant  $\alpha$  and the persistence parameter  $\rho$  at different quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  for different subsamples. The gray areas indicate 95% bootstrapped confidence bands. The subsamples are chosen based on the estimated break points of the DQ-test for CPI inflation. For PCE inflation we can only plot estimates for the last two regimes as the data series starts later.

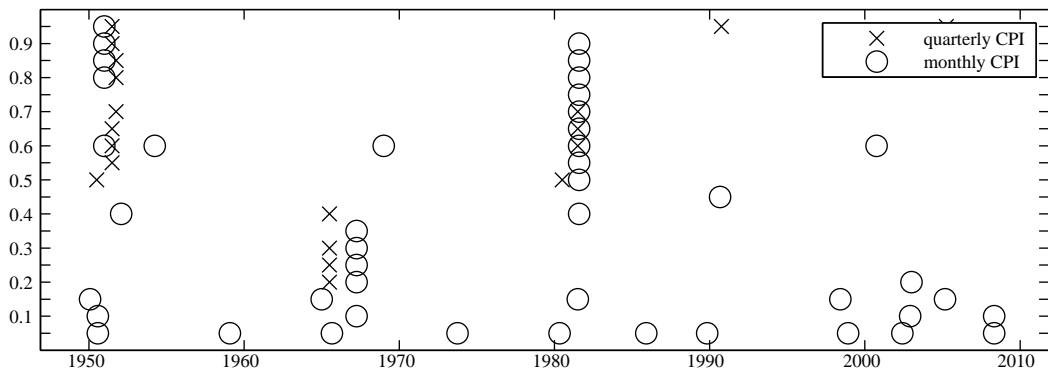


Figure 12: Estimated break points at different quantiles for CPI inflation after controlling for a mean shift. Notes: The graph shows breakpoints estimated with the  $SQ_\tau$ -test at all quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  (vertical axis) for CPI inflation demeaned separately for the subsamples 1947Q2-1951Q3, 1951Q3-1981Q3 and 1981Q3-2011Q2.

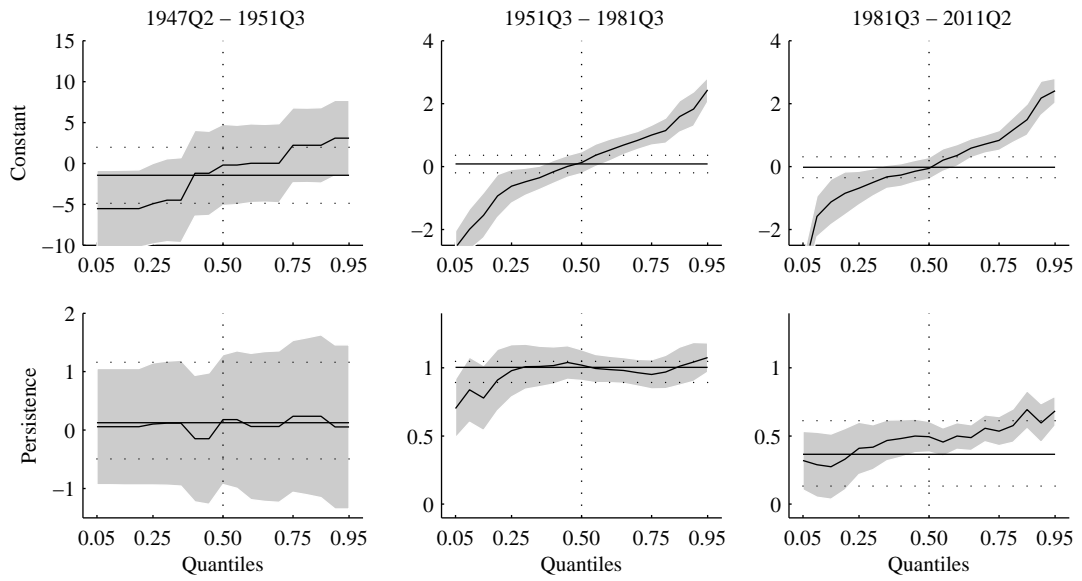


Figure 13: Inflation persistence (quarterly CPI) after controlling for a mean shift. Notes: The graphs show estimates of the constant  $\alpha$  and the persistence parameter  $\rho$  at different quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  for different subsamples. The gray areas indicate 95% bootstrapped confidence bands. The subsamples are chosen based on the estimated break points of the DQ-test for CPI inflation. CPI inflation has been demeaned separately for the subsamples. The horizontal line shows Hansen's (1999) median unbiased estimator together with 95% grid-t bootstrap confidence bands.

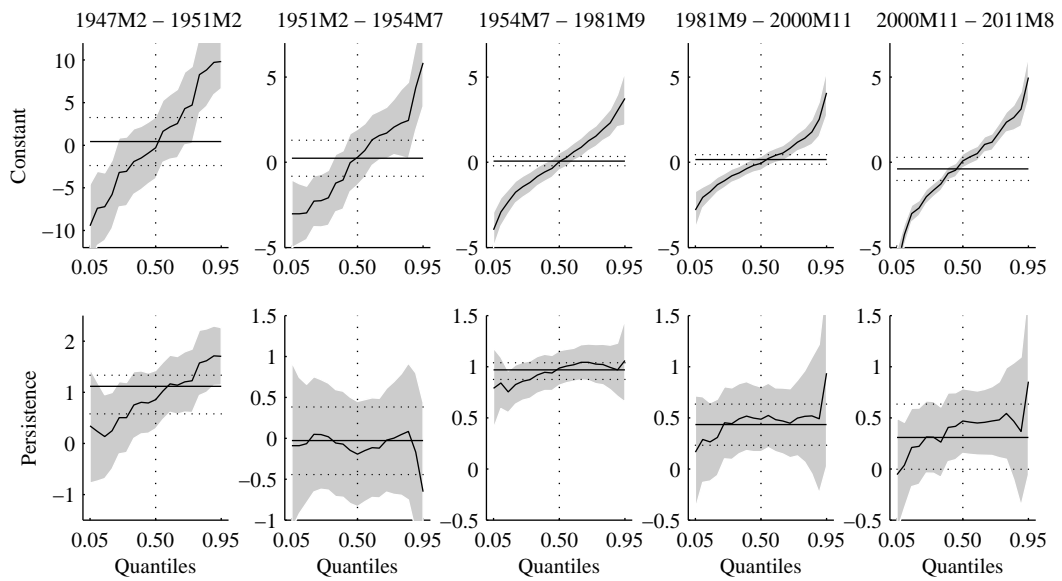


Figure 14: Inflation persistence (monthly CPI) after controlling for a mean shift. Notes: The graphs show estimates of the constant  $\alpha$  and the persistence parameter  $\rho$  at different quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  for different subsamples. The gray areas indicate 95% bootstrapped confidence bands. The subsamples are chosen based on the estimated break points of the DQ-test for CPI inflation. CPI inflation has been demeaned separately for the subsamples. The horizontal line shows Hansen's (1999) median unbiased estimator together with 95% grid-t bootstrap confidence bands.

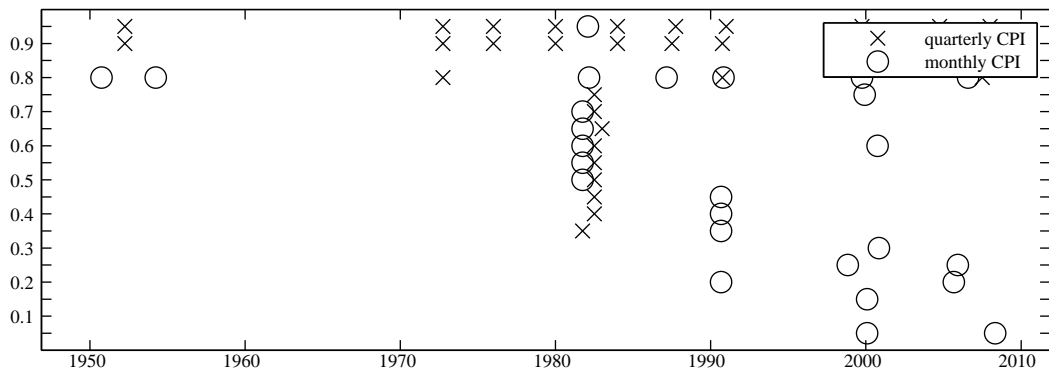


Figure 15: Estimated break points at different quantiles for CPI inflation after controlling for mean and variance shifts. Notes: The graph shows breakpoints estimated with the  $SQ_\tau$ -test at all quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  (vertical axis) for CPI inflation normalized separately for the subsamples 1947Q2-1951Q3, 1951Q3-1981Q3 and 1981Q3-2011Q2.



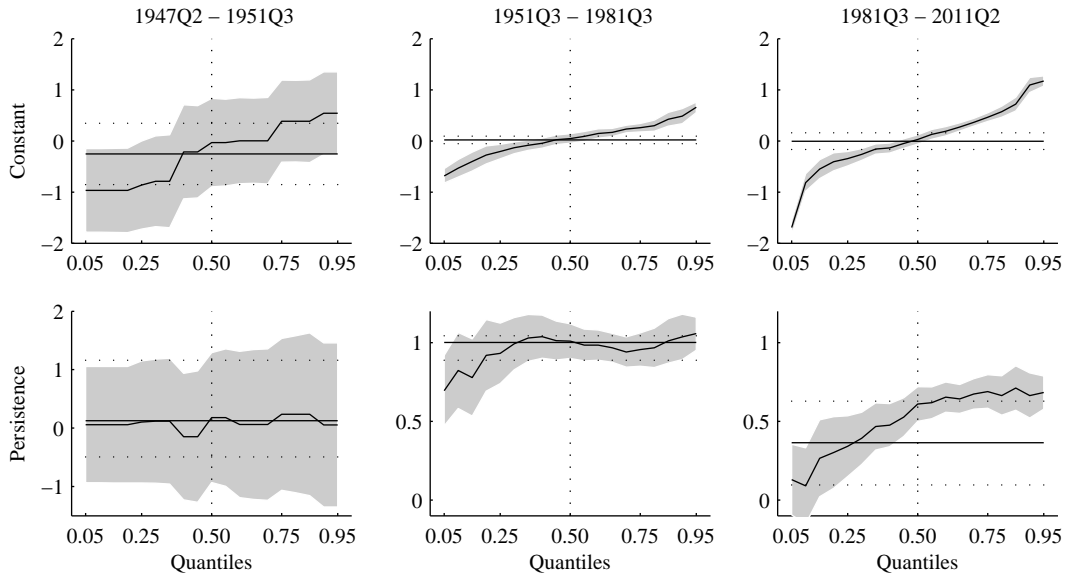


Figure 16: Inflation persistence (quarterly CPI) after controlling for mean and variance shifts. Notes: The graphs show estimates of the constant  $\alpha$  and the persistence parameter  $\rho$  at different quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  for different subsamples. The gray areas indicate 95% bootstrapped confidence bands. The subsamples are chosen based on the estimated break points of the baseline DQ-test for CPI inflation. CPI inflation has been normalized separately for the subsamples. The horizontal line shows Hansen's (1999) median unbiased estimator together with 95% grid-t bootstrap confidence bands.

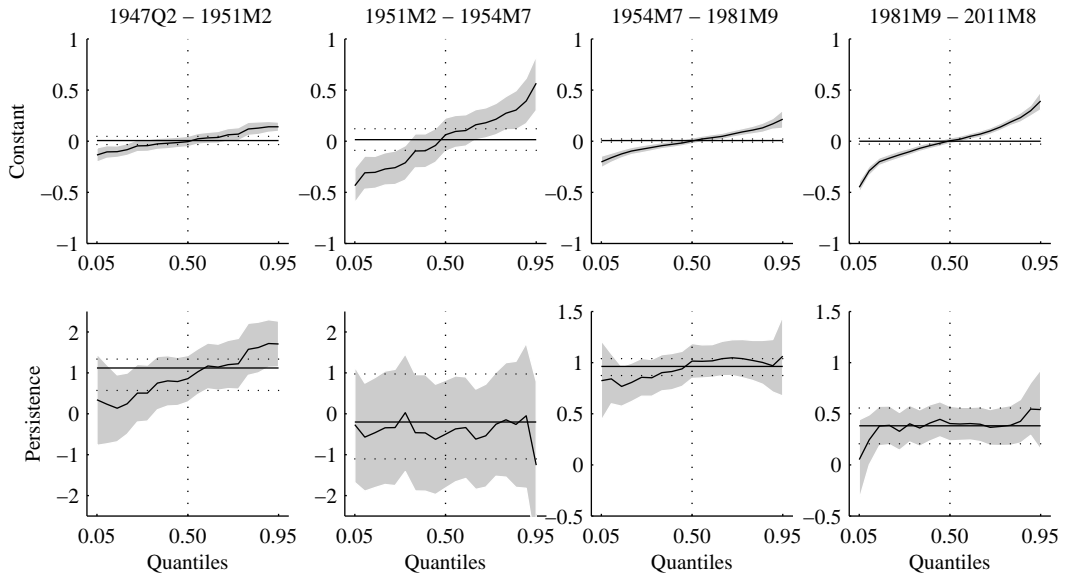


Figure 17: Inflation persistence (monthly CPI) after controlling for mean and variance shifts. Notes: The graphs show estimates of the constant  $\alpha$  and the persistence parameter  $\rho$  at different quantiles  $\tau \in \{0.05, 0.1, \dots, 0.95\}$  for different subsamples. The gray areas indicate 95% bootstrapped confidence bands. The subsamples are chosen based on the estimated break points of the baseline DQ-test for CPI inflation. CPI inflation has been normalized separately for the subsamples. The horizontal line shows Hansen's (1999) median unbiased estimator together with 95% grid-t bootstrap confidence bands.