

Lein, Sarah M.; Maag, Thomas

**Working Paper**

## The formation of inflation perceptions: Some empirical facts for European countries

KOF Working Papers, No. 204

**Provided in Cooperation with:**

KOF Swiss Economic Institute, ETH Zurich

*Suggested Citation:* Lein, Sarah M.; Maag, Thomas (2008) : The formation of inflation perceptions: Some empirical facts for European countries, KOF Working Papers, No. 204, ETH Zurich, KOF Swiss Economic Institute, Zurich, <https://doi.org/10.3929/ethz-a-005666792>

This Version is available at:

<https://hdl.handle.net/10419/50341>

**Standard-Nutzungsbedingungen:**

Die Dokumente auf EconStor dürfen zu eigenen wissenschaftlichen Zwecken und zum Privatgebrauch gespeichert und kopiert werden.

Sie dürfen die Dokumente nicht für öffentliche oder kommerzielle Zwecke vervielfältigen, öffentlich ausstellen, öffentlich zugänglich machen, vertreiben oder anderweitig nutzen.

Sofern die Verfasser die Dokumente unter Open-Content-Lizenzen (insbesondere CC-Lizenzen) zur Verfügung gestellt haben sollten, gelten abweichend von diesen Nutzungsbedingungen die in der dort genannten Lizenz gewährten Nutzungsrechte.

**Terms of use:**

*Documents in EconStor may be saved and copied for your personal and scholarly purposes.*

*You are not to copy documents for public or commercial purposes, to exhibit the documents publicly, to make them publicly available on the internet, or to distribute or otherwise use the documents in public.*

*If the documents have been made available under an Open Content Licence (especially Creative Commons Licences), you may exercise further usage rights as specified in the indicated licence.*

# KOF Working Papers

The Formation of Inflation Perceptions – Some Empirical Facts  
for European Countries

Sarah M. Lein and Thomas Maag



# KOF

ETH Zurich  
KOF Swiss Economic Institute  
WEH D 4  
Weinbergstrasse 35  
8092 Zurich  
Switzerland

Phone +41 44 632 42 39  
Fax +41 44 632 12 18  
[www.kof.ethz.ch](http://www.kof.ethz.ch)  
[kof@kof.ethz.ch](mailto:kof@kof.ethz.ch)

# The Formation of Inflation Perceptions: Some Empirical Facts for European Countries\*

Sarah M. Lein<sup>†</sup> and Thomas Maag<sup>‡</sup>

KOF Working Paper No. 204, September 2008

This Version: December 2009

## Abstract

This paper investigates how households form their perceptions of consumer price inflation. Using data from the harmonized EU consumer survey, we find that inflation perceptions are inefficient and highly heterogeneous, yet contemporaneously related to the actual rate of inflation. Consequently, we estimate how often households update their beliefs employing Carroll's (2003) epidemiological model. Our results indicate that inflation perceptions are generally less responsive to new information than expectations. Unlike studies on expectations, we cannot confirm that a constant fraction of the population updates information every month. Also, the cross-sectional heterogeneity of perceptions is higher than implied by infrequent updating alone.

*JEL classification:* E31, E50, D83.

*Keywords:* inflation perceptions, belief formation, sticky information models.

---

\*We thank Helge Berger, Michael Lamla, Jan-Egbert Sturm as well as participants of the 23rd Annual Congress of the European Economic Association, Milan, Italy, and the 4th Macroeconomic Research Meeting, University of Bonn, Germany, for helpful comments and suggestions. Financial support by the Swiss National Science Foundation (SNF) is gratefully acknowledged. The views expressed in this paper reflect those of the authors and not necessarily those of the Swiss National Bank or the ETH Zurich.

<sup>†</sup>Swiss National Bank, CH-8022 Zurich, Switzerland, E-mail: sarah.lein@snb.ch

<sup>‡</sup>Corresponding author. KOF Swiss Economic Institute, ETH Zurich, CH-8092 Zurich, Switzerland, E-mail: maag@kof.ethz.ch

# 1 Introduction

This paper investigates how households form inflation perceptions, defined as the beliefs at time  $t$  about the actual rate of consumer price inflation between month  $t-12$  and  $t$ . Economic theory suggests that expectations about future inflation have predominant implications for investment, saving and consumption decisions. We argue, however, that it is just as important to investigate perceptions of current annual inflation for two main reasons. First, inflation perceptions are an important determinant of inflation expectations. This is suggested by results of Benford and Driver (2008) who investigate data from a special issue of the Bank of England Inflation Attitudes Survey that asks households about how they form their inflation expectations. Benford and Driver (2008) document that more than 40% of households consider their perception of current inflation to be a very important factor in expectation formation. Inflation perceptions are more important than the other factors mentioned, which include interest rates, the central bank policy target and media reports. Second, perceptions allow for better tests of rationality and models of belief formation than expectations. Working with inflation perceptions, the benchmark for the belief that a rational household should adopt is relatively well-defined. It is the publicly available, official rate of inflation.

The empirical literature on inflation perceptions is scant, both in absolute terms and relative to the literature on inflation expectations. Only recently, the rise in inflation perceptions coinciding with the euro cash changeover in the European Monetary Union has drawn increased research attention.<sup>1</sup> Abstracting from the euro cash changeover, an earlier literature comprises a small number of papers that investigate household-level data. Using quantitative survey data from the U.S., Bryan and Venkatu (2001a, 2001b) find that inflation perceptions of households are significantly biased. Furthermore, they report that the accuracy of inflation perceptions correlates with demographic characteristics. Jonung (1981) and Palmqvist and Strömberg (2004) document similar patterns using survey data from Sweden.

The goal of our paper is to understand how inflation perceptions of households are related to the actual rate of inflation in a sample of 12 European countries. We aim to provide general evidence rather than focusing on specific factors associated with the euro cash changeover or

---

<sup>1</sup>The deviation of perceived from actual inflation rates is documented in ECB (2005). The literature investigates several explanations for this rise, including increased information processing requirements due to conversion rates, overreaction to prices of frequently bought items and anchoring of perceptions to prior beliefs. See Ehrmann (2006), Aucremanne, Collin, and Stragier (2007), Doehring and Mordonu (2007), Dziuda and Mastrobuoni (2006), Aalto-Setälä (2006), Fluch and Stix (2007) and references therein.

with socioeconomic characteristics. Our analysis begins by presenting evidence on the dynamics and rationality of inflation perceptions. We find that inflation perceptions fail rationality tests and that perceptions exhibit a high degree of cross-sectional heterogeneity. These broad patterns are consistent with the epidemiological model of belief formation proposed by Carroll (2003). In this model, only a fraction of households update their beliefs with the latest information in each period. The rest of households is assumed to stick to outdated beliefs. This model implies an inertial response of the population mean of perceived inflation to changes in the actual rate of inflation. We formally test whether the dynamics of the survey mean and the cross-sectional heterogeneity of inflation perceptions can be explained by the epidemiological model. This is, to some extent, also an assessment of the sticky information hypothesis put forward by Mankiw and Reis (2002).<sup>2</sup>

We find that a share of around 11% of consumers in the euro area update their inflation perceptions within a quarter of a year. These estimates are lower than the updating frequencies reported by studies relying on survey data about inflation expectations. For European countries, Döpke, Dovern, Fritsche and Slacalek (2008a, 2008b) find that between 20 and 30 percent of households (and firms) update their expectations within a given quarter.<sup>3</sup> However, we find that the epidemiological model does not adequately describe perception formation in our sample of European countries. In particular, we show that the cross-sectional heterogeneity of inflation perceptions is much higher than predicted by the model. We therefore conclude that the transmission of information to households and the formation of beliefs should be described by approaches that include alternative mechanisms which generate an extra degree of cross-sectional heterogeneity of survey responses.

The paper is structured as follows. Section 2 discusses models of perception formation, including the rational perceptions and the epidemiological perceptions model. Section 3 presents the dataset which is based on the Joint Harmonized EU Consumer Survey and the Swedish Consumer Tendency Survey. Section 4 investigates general properties of inflation perceptions and tests the rational perceptions hypothesis. Section 5 assesses the epidemiological model of perception formation. Section 6 concludes.

---

<sup>2</sup>Similar to the epidemiological model, the sticky information model of Mankiw and Reis (2002) assumes that agents that do not update their information sets stick to outdated beliefs. But unlike in the epidemiological model, agents continue to dynamically compute beliefs based on their outdated information sets. The epidemiological model assumes that agents stick to their static belief formed at the time of the last update.

<sup>3</sup>For the U.S., Carroll (2003) and Khan and Zhu (2006) estimate updating frequencies that lie in the same range.

## 2 Models of Perception Formation

We initially assume that households form rational beliefs about actual consumer price inflation. Following the rational expectations literature, we define perceptions  $\pi_t^p$  of actual inflation  $\pi_t$  to be rational if:

$$\pi_t^p = E_t \pi_t$$

where  $E_t$  is the expectation of actual inflation conditional on the public information set  $\Omega_t$  available at time  $t$ . The rational perceptions hypothesis assumes that agents employ all available information to form beliefs about actual inflation. This hypothesis can be tested by investigating whether inflation perceptions are unbiased and information efficient.<sup>4</sup>

As an alternative to rational perceptions, we consider the epidemiological model of belief formation proposed by Carroll (2003). The epidemiological model is based on the idea that households form inflation expectations by probabilistically acquiring new information from media reports. Carroll (2003) assumes that media reports transmit expectations of professional forecasters which are subsequently adopted by households. Every household has a constant probability  $\lambda$  of encountering media reports and absorbing the most recent professional forecasts in a given month  $t$ . This assumption implies that the population mean of inflation expectations can be written as a partial adjustment model:

$$\pi_{t,t+12}^e = \lambda \pi_{t,t+12}^m + (1 - \lambda) \pi_{t-1,t+11}^e$$

where  $\pi_{t,t+12}^e$  is the cross-sectional mean of household expectations about the one year ahead inflation rate,  $\pi_{t,t+12}^m$  is the inflation forecast of professional forecasters transmitted by the media and  $\pi_{t-1,t+11}^e$  is the expected one year ahead inflation rate of households in the previous month. The coefficient  $\lambda$  is equal to the proportion of households that update their inflation expectations with the new expectation of professional forecasters in a given month.<sup>5</sup> The

---

<sup>4</sup>If  $\Omega_t$  contains all information including  $\pi_t$ , perceptions are rational if the identity  $\pi_t^p = \pi_t$  holds. Still, we rely on less restrictive tests of unbiasedness and information efficiency to assess the rational perceptions hypothesis since the timing of the household survey is not identical for all households. As will be discussed in Section 3, the household survey is conducted during the first three weeks of each month.

<sup>5</sup>Carroll (2003) and Döpke, Doornik, Fritsche, and Slacalek (2008b) estimate this model for the U.S. and Europe respectively. They find that consumers in the U.S. update their information about once a year, in Europe about once in eighteen months.

epidemiological model is related to the concept of sticky information introduced by Mankiw and Reis (2002, 2006). The main assumption in sticky information models is that in each period, only a fraction of agents acquire new information about the state of the economy to compute a new path of optimal behavior. Those agents who update are assumed to form rational expectations based on  $\Omega_t$ . Consequently, new information disperses slowly throughout the population and has a gradual and delayed effect on the aggregate behavior of agents.

The critical decision that has to be made when testing models of expectation formation concerns the identification of the new information that agents use to update their beliefs. It is inherently difficult to identify the ex-ante rational value of expectations. The literature on expectation formation mainly employs two benchmark measures: the actual rate of inflation materialized in 12 months and inflation expectations of professional forecasters, as in the model of Carroll (2003). Both benchmark measures of rational expectations can be criticized on theoretical grounds.

Using the materialized 12 months ahead inflation rate may be flawed if agents assign positive probability to an important event that does not materialize. In retrospect, one will then observe biased expectations and autocorrelated expectations errors even if expectation formation was rational.<sup>6</sup> Similarly, even rational agents may not identify a regime change as being permanent at first sight, in which case expectations would again fail rationality tests. Andolfatto, Hendry, and Moran (2008) confirm this argument in a simulation study building on the rational expectations New Keynesian model. The model assumes that agents face a signal extraction problem as they only have incomplete information about the changing inflation target of the monetary authority. Calibrating the model to fit U.S. business cycle statistics, Andolfatto, Hendry, and Moran (2008) show that conventional tests of rational expectations incorrectly reject rationality in about 30% of the simulated samples that span 80 quarters.

Using expectations of professional forecasters as the rational benchmark has some weaknesses, too. Several studies report that professional forecasts are biased, see, e.g., Ang, Bekaert, and Wei (2007), Mehra (2002) and Thomas (1999). A rational household that is aware of this might thus not rely on professional forecasts when forming inflation expectations. Moreover, professional forecasters usually disagree. It is unclear which forecast households

---

<sup>6</sup>This potential pitfall is also known as the “Peso Problem”, see Jonung and Laidler (1988). For these reasons it is common practice to assess rationality of expectations only over long time periods.



will refer to. In particular, the literature that investigates strategic forecasting commonly assumes that forecasts far off the consensus catch more media attention.<sup>7</sup> Therefore, the central tendency of professional forecasts might not correspond to the inflation forecast that households observe in the media and use to update their beliefs with.

Unlike for inflation expectations, the rational benchmark seems well-defined for inflation perceptions: It is the actual rate of consumer price inflation as published by national statistical offices. Actual consumer price inflation is, with a short publication lag, unambiguously available to all households, be it as an official news releases of the national authority, be it through media reports. Therefore, in line with the reasoning of Jonung and Laidler (1988), inflation perceptions might be better suited to assess rationality and models of belief formation than inflation expectations.

We thus rewrite the epidemiological model of Carroll (2003) to a partial adjustment model in which consumers update their inflation perceptions with the actual rate of inflation. We consider two versions of the model. The first specification assumes that households update using the contemporaneous, actual rate of inflation. Since the official inflation rate is published rather in the beginning to the middle of the following month, the model assumes that households compile new information based on price changes they observe during economic interactions in the current month. Consumers that do not update their information set stick to the same inflation perception as in the previous month. The resulting partial adjustment model (1) can be written as follows:

$$\pi_t^p = \lambda\pi_t + (1 - \lambda)\pi_{t-1}^p \quad (1)$$

The second specification assumes that households use the most recent available official inflation figure to update their information sets. Due to the publication lag of the official inflation figure, we therefore test whether consumers absorb the one month lagged inflation rate. The partial adjustment model (2) is given by:

$$\pi_t^p = \lambda\pi_{t-1} + (1 - \lambda)\pi_{t-1}^p \quad (2)$$

---

<sup>7</sup>Making biased forecasts far from the average might in turn be rational behavior by professional economists, see e.g. Ehrbeck and Waldmann (1996) and Laster, Bennett, and Geoum (1999). In the model of Laster, Bennett, and Geoum (1999), forecasters are not only paid on basis of the accuracy of their forecast but also on basis of the media attention they are able to catch. The latter can be obtained by deviating significantly from the average.

Similar to Carroll (2003), this model implicitly assumes that consumers update their beliefs with the latest inflation rate published in the media.

### 3 Data

In the European Union, household inflation perceptions are surveyed as part of the Joint Harmonized European Union Consumer Survey.<sup>8</sup> In each member state, national institutes survey about 1,500 households during the first three weeks of every month. In July 2007, the overall sample covers 39,900 consumers in 27 member states. Inflation perceptions are captured by asking households: “How do you think that consumer prices have developed over the last 12 months? They have...”. Respondents are asked to indicate their beliefs on an ordinal scale with five response categories given by: “Risen a lot, risen moderately, risen slightly, stayed about the same, fallen”.

In line with recent literature on expected inflation we quantify the qualitative response data using the probability method.<sup>9</sup> We employ the probability method for 5-category response data as proposed by Batchelor and Orr (1988). Inflation perceptions are identified by assuming that perceptions are unbiased with respect to actual consumer price inflation during the sample period. Under this identification scheme, the probability method allows to compute the mean and standard deviation of inflation perceptions among respondents in a given month. A detailed description of the approach and its identifying assumptions are provided in Appendix A.2.<sup>10</sup>

To assess the robustness of the quantification method, we additionally discuss estimations that are based on direct quantitative survey data obtained from the Swedish Consumer Tendency Survey. This survey has been capturing quantitative inflation perceptions on a monthly

---

<sup>8</sup>The consumer survey consists of 15 qualitative questions that capture the financial situation, perceived economic conditions and planned savings and spending. This standard questionnaire is translated into national languages and may include additional country specific questions, see European Commission (2007).

<sup>9</sup>Recent contributions that use the probability method to quantify expected inflation are Berk (1999) and Forsells and Kenny (2004) who quantify EU consumer survey data and Mankiw, Reis, and Wolfers (2004) who quantify qualitative response data from the University of Michigan Survey of Consumers.

<sup>10</sup>Batchelor and Orr (1988) extend the probability method developed in Theil (1952) and Carlson and Parkin (1975) to 5-category response data. Perceived inflation and implied standard deviation of perceptions are given by equations (A.3) and (A.4) in the Appendix. Maag (2009) assesses the empirical performance of the probability method using data both on qualitative and quantitative inflation perceptions taken from the Swedish Consumer Tendency Survey. It is found that in Swedish data, the method applied in this paper generates series that have a correlation of 0.97 (0.86 in first differences) with actual quantitative inflation perceptions during 1996–2008.

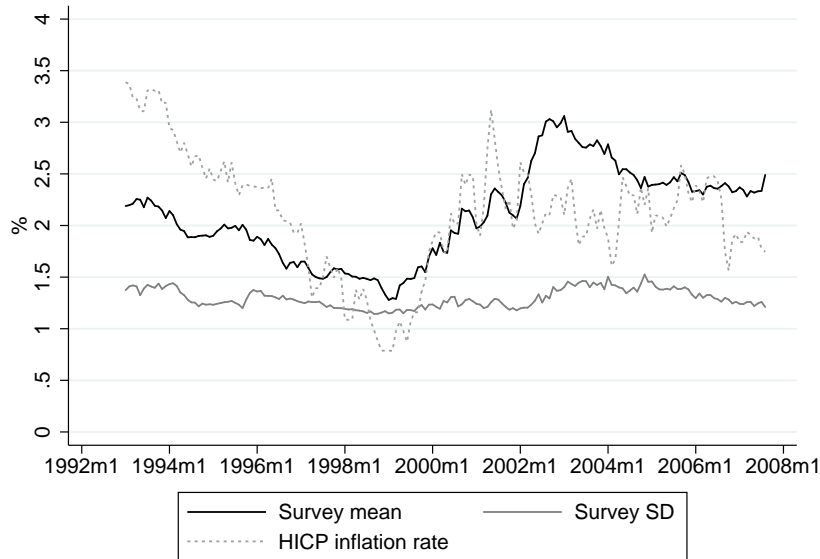


Figure 1: Mean and standard deviation of inflation perceptions in the euro area

*Notes:* The figure shows quantified mean and cross-sectional standard deviation of inflation perceptions in the euro area. Qualitative response data is quantified using the 5-category probability method under the assumption that perceptions are unbiased.

basis since January 1996.<sup>11</sup>

As a measure for actual inflation we use the annual percent change in the Harmonized Index of Consumer Prices (HICP) as published by Eurostat. The actual annual inflation rate is given by  $\pi_t = 100 \left( \frac{P_t}{P_{t-12}} - 1 \right)$ , where  $P_t$  is the level of the HICP index at the end of month  $t$ . We assume that rational individuals absorb this inflation rate as their belief about actual inflation.<sup>12</sup> For testing the rationality of households, we also compute a price index for out-of-pocket expenditures. The out-of-pocket expenditures index covers non-durable goods and consumer services that are frequently purchased and paid in cash. It is based on disaggregate HICP data obtained from Eurostat.<sup>13</sup>

We consider a sample of 10 euro area countries, Sweden, United Kingdom and the euro area (EA) aggregate.<sup>14</sup> For most countries, the sample includes 176 monthly observations

<sup>11</sup>See GfK (2002) for a description of the Swedish Consumer Tendency Survey.

<sup>12</sup>We have employed real time data for HICP inflation rate where available and tested the models. Differences to using ex post data were insignificant. This result is not surprising, as revisions in inflation rates are rather rare and small. Therefore, we report the ex post data results, as we have more data points available for earlier years in the time series dimension.

<sup>13</sup>We compute the out-of-pocket expenditures index as a consumption weighted average of price series on COICOP 2 to 4 digit level. The included items are food, beverages, tobacco, non-durable household goods, transport services, fuel, postal services, hotels, restaurants and hairdressing.

<sup>14</sup>Of the 12 countries that introduced euro banknotes and coins in 2002, Luxembourg and Portugal are

spanning 01/1993 to 08/2007. It may be reduced depending on the joint availability of survey data and HICP inflation rates. Table A.1 in the Appendix provides an overview. To account for potential structural breaks coinciding with the euro cash changeover, we provide additional estimation results for the subperiods 01/1993–12/2001 and 01/2003–08/2007.

Figure 1 shows perceived inflation in the euro area. The mean of inflation perceptions clearly exhibits the so called euro cash changeover effect. Perceived inflation rises above actual inflation in 2002 and peaks at about 3% in 01/2003. The figure indicates that the gap closes in early 2004. Perceived inflation has a low standard deviation of 0.44%, while actual HICP inflation has a standard deviation of 0.58% during 1993–2007. Relative to the variability of inflation, the quantified cross-sectional standard deviation of inflation perceptions seems high. It averages at 1.29% during 1993–2007. As will be discussed below, this quantified series even tends to underestimate the actual heterogeneity of perceptions.

For the further analysis, stationarity properties of perceived and actual inflation are critical. We discuss unit root properties and cointegration of actual and perceived inflation both on a country-by-country basis and in a panel framework. Our analysis begins by testing the null-hypothesis of a unit root using the augmented Dickey-Fuller (ADF) test. We apply the sequential model selection procedure following Perron (1988). Additionally, the null hypothesis of stationarity is tested using the Kwiatkowski, Phillips, Schmidt, and Shin (1992) (KPSS) test. Tables A.2 and A.3 in the Appendix summarize the results for HICP inflation and perceived inflation, respectively. The ADF-tests indicate that actual inflation is stationary in 4 out of 12 countries, whereas perceived inflation is stationary in 2 countries. The KPSS test always rejects its null hypothesis of stationarity. Clearly, all first differences are stationary. As perceived inflation rates have substantially increased during the euro cash changeover in most countries of the euro area, it might well be the case that the ADF tests fail to reject the null hypothesis due to a level shift in the underlying series. We therefore additionally apply a unit root test that allows for a deterministic level shift following Saikkonen and Lütkepohl (2002) and Lanne, Lütkepohl, and Saikkonen (2002). The results are reported in Table A.4 in the Appendix and the conclusions are unchanged. We cannot reject the null of a unit root in perceived and actual inflation in most countries.

---

not included since no survey data on inflation perceptions is available. Aggregate actual HICP inflation and perceived inflation are computed as weighted means of the euro area series. Weights are given by private domestic consumption expenditures.

While the sample includes a reasonable number of monthly observations, it covers a relatively short time span of only 14 years. As Pierse and Snell (1995) show, the power of a unit root test primarily depends on the time span rather than on the sampling frequency. To improve power, we thus additionally investigate unit root properties in a panel setting using the Im, Pesaran, and Shin (2003) (IPS) and Pesaran (2007) (CIPS) panel unit root tests. These tests assesses the null hypothesis of a unit root in all countries against the alternative that inflation is stationary in a significant number of countries. Both test allow for heterogeneous short run dynamics and deterministic terms. An important restriction of the IPS test is that dependence of inflation across countries is only accounted for by cross-sectional demeaning. Other forms of cross-sectional dependence that cannot be captured by a homogenous common time effect may induce a positive size bias (Breitung and Pesaran, 2005). In our panel of 12 interconnected European economies, overrejection is a relevant issue.<sup>15</sup> The CIPS test proposed by Pesaran (2007) is a more reliable alternative. This test allows for residual correlation that is generated by a heterogeneous single-factor structure. Its limitation remains the assumption of a single common factor. Contrary to the country-by-country tests, the panel unit root tests shown in Table A.5 clearly reject the null hypothesis of a unit root in inflation. For higher lag orders, the CIPS test clearly rejects the null hypothesis of a unit root in all series, both in the period 1993–2007 period and in subperiod 1993–2001.<sup>16</sup> The null of a unit root is not rejected in the shorter subperiod 2003–2007. Perceived inflation appears to be more persistent than actual inflation. Both tests cannot reject the null hypothesis in the samples 1993–2007 and 1993–2001.

Due to the ambiguous findings on the stationarity properties, we additionally investigate cointegration of actual and perceived inflation. Obviously, one would expect that perceived and actual inflation move together proportionally in the long run. Hence, if the series are  $I(1)$  they should be cointegrated. Table A.6 in the Appendix reports results from Johansen trace tests on the cointegration rank. The table shows trace statistics for the null hypothesis of no cointegration ( $r = 0$ ) and the null hypothesis of one cointegration relation ( $r = 1$ ) between actual and perceived inflation during 1993–2007. The tests indicate that actual and perceived

---

<sup>15</sup>Even after controlling for common time effects, the absolute cross-sectional correlation coefficients still average at 0.32. See Wang and Wen (2007) for a discussion of potential sources of international synchronization in inflation rates.

<sup>16</sup>We consider the specifications that include 3 or 4 lags since under-fitting can lead to considerable size distortions as shown by Im, Pesaran, and Shin (2003).

inflation are cointegrated only in Greece and Ireland. Consistent with the ADF test, the full rank results indicate that an estimation in levels is appropriate for Finland and Sweden. The general picture of no cointegration remains unaltered in the subperiods 1993–2001 and 2003–2007, as Tables A.7 and A.8 show.

To gain statistical power we also employ panel cointegration tests. We use the residual based tests proposed by Pedroni (1999, 2004). The null hypothesis of no cointegration in all countries is tested against the alternative hypothesis that a cointegration relation exists in a significant number of countries. Cointegration coefficients and short run dynamics are allowed to differ across countries.<sup>17</sup> Table A.9 in the Appendix reports the parametric (analogue to the augmented Dickey-Fuller  $t$ -statistic) and nonparametric (analogue to the Phillips and Perron  $t$ -statistic) panel and group mean  $t$ -statistics. The panel tests indicate that perceived and actual inflation are cointegrated over the full sample as well as in the two subperiods.<sup>18</sup> Although the panel result that actual and perceived inflation are cointegrated is intuitively appealing, the result might also mirror that actual inflation and inflation perceptions are stationary. The panel unit root tests generally reject the null hypothesis of a unit root in actual inflation and, on a country-by-country basis, a cointegration relation is detected only in a small number of countries.

The ambiguous results are in line with the mixed findings of the empirical literature. Surveying this literature, Altissimo, Ehrmann, and Smets (2006) conclude that empirical work is rather in favor of stationarity of euro area consumer price inflation.<sup>19</sup> Moreover, from a theoretical economic viewpoint it seems reasonable to assume that inflation and inflation perceptions are no unit root processes in the sample period considered here. We therefore estimate our baseline specifications in levels. To assess the robustness, we additionally provide estimation results in first differences.

---

<sup>17</sup>The limitation of this test is that it does not account for cross-sectional relations that cannot be removed by simple cross-section demeaning. Similar to the panel unit root tests, this may lead to size distortions (Banerjee, Marcellino and Osbat, 2004).

<sup>18</sup>This result is confirmed by the five other test statistics proposed by Pedroni (1999, 2004).

<sup>19</sup>Still, some recent studies cannot reject the null hypothesis of a unit root in inflation, see e.g. O'Reilly and Whelan (2005).

## 4 Explorative Analysis

### 4.1 Accuracy of Inflation Perceptions

This section highlights some general statistical properties pertaining to the dynamics of inflation perceptions and empirically motivates the epidemiological model of belief formation. The analysis begins by investigating the accuracy of inflation perceptions. As Table 1 shows, the accuracy with respect to actual HICP inflation varies quite substantially between countries. The mean absolute error (MAE) ranges between 0.48 and 1.72 percent. It averages at 0.86% during 1993–2007. This seems relatively high, given that the quantified inflation perceptions are unbiased by assumption. Compared to the accuracy of inflation expectations as documented in the literature, inflation perceptions are only slightly more accurate.<sup>20</sup> This is particularly noteworthy, as inference about current and past inflation entails substantially less uncertainty compared to inference about the 12 months ahead inflation.

Table 1 additionally reports the correlation ( $\rho$ ) of perceived with actual HICP inflation. Looking at the euro area aggregate, this correlation drops from 0.90 in the period 1993–2001 to -0.04 in the period 2003–2007. The observed decline in correlation is broadly consistent across euro area countries. In Sweden and the United Kingdom, correlations are stable or even increase over time. The table also shows correlation of perceived inflation with one month lagged actual HICP inflation and with inflation in the out-of-pocket expenditures index. Correlations with one month lagged inflation are virtually identical to contemporaneous correlations. No clear pattern emerges regarding the correlation between perceived inflation and out-of-pocket inflation. For the euro area aggregate, the correlation is higher than the correlation of perceived with actual HICP inflation. At country level, however, the correlation is higher only in Belgium, Germany, Spain and Italy.

### 4.2 Rationality of Inflation Perceptions

A large literature investigates rationality of inflation expectations. We borrow rationality tests from this literature to assess inflation perceptions. Along the lines of Jonung and Laidler (1988), it may be argued that inflation perceptions are more adequate than inflation

---

<sup>20</sup>E.g., Mankiw, Reis, and Wolfers (2004) report that 12 months ahead inflation expectations taken from the University of Michigan Survey of Consumers have a MAE of 1.07% and a RMSE of 0.85% in the period 1982–2002.

Table 1: Accuracy of inflation perceptions by country

<i>Country</i>	1993–2007					93–01	03–07
	MAE	RMSE	$\rho$	$\rho_{lag}$	$\rho_{OOP}$	$\rho$	$\rho$
AT Austria	0.74	0.70	0.47	0.48	0.38	0.49	0.19
BE Belgium	0.71	0.60	0.47	0.50	0.52	0.72	0.40
DE Germany	0.69	0.58	0.23	0.26	0.48	0.78	-0.58
EA Euro Area	0.76	0.72	0.45	0.45	0.77	0.90	-0.04
EL Greece	1.72	3.76	-0.17	-0.18	n.a.	0.74	-0.34
ES Spain	0.93	1.06	0.32	0.33	0.68	0.78	0.32
FI Finland	0.86	0.90	0.31	0.35	0.22	0.73	0.57
FR France	0.69	0.57	0.50	0.51	0.24	0.55	0.26
IE Ireland	0.84	0.91	0.71	0.74	0.58	0.90	0.61
IT Italy	0.95	1.11	0.35	0.34	0.69	0.83	0.78
NL Netherlands	0.84	1.01	0.42	0.46	0.37	0.87	0.58
SE Sweden	0.66	0.56	0.72	0.70	0.42	0.77	0.73
SEq Sweden quant.	0.48	0.61	0.66	0.65	0.36	0.70	0.68
UK United Kingdom	0.63	0.51	0.35	0.35	0.29	-0.02	0.61

*Notes:* The table shows mean absolute error (MAE), root mean squared error (RMSE) and correlation ( $\rho$ ) of perceived inflation relative to actual HICP inflation.  $\rho_{lag}$  and  $\rho_{OOP}$  denote correlation coefficients of perceived inflation with one month lagged HICP inflation and out-of-pocket expenditures inflation, respectively.

expectations for testing rationality of households, since the rational benchmark is relatively well-defined. Tests of rationality include the related aspects of unbiasedness and information efficiency. Since the quantification method imposes unbiasedness, we can only assess bias in the quantitative data for Sweden. During 01/1996–08/2007, inflation perceptions of Swedish households are unbiased with a statistically insignificant average perception error of  $\pi_t - \pi_t^p = 0.01\%$ . This finding is in contrast to significant biases documented by Bryan and Venkatu (2001a, 2001b) for U.S. survey data.<sup>21</sup>

We thus assess rationality by testing whether agents efficiently use available information to form their perceptions of inflation. As a first test of efficiency, we assess whether perception errors, defined as the difference between actual and perceived inflation, are serially correlated. Clearly, if perception formation is rational, past perception errors have no predictive content for subsequent errors. The first panel of Table 2 reports estimation results on the serial correlation of perception errors over non-overlapping periods. Both in the euro area and in Sweden, perception errors exhibit pronounced serial correlation.<sup>22</sup> The results for

<sup>21</sup>Using monthly household survey data of the Federal Reserve Bank of Cleveland, Bryan and Venkatu (2001a, 2001b) show that inflation perceptions (and expectations) of U.S. households average several percentage points above actual consumer price inflation.

<sup>22</sup>We only report results for the EA and Sweden, the remaining country-by-country results are broadly in



Sweden using quantified and quantitative (denoted by the country code *SEq*) survey data are consistent.

The second panel of Table 2 investigates whether perceptions efficiently incorporate publicly available information. In defining the relevant information set we follow the literature on inflation expectations and include the money market rate and the unemployment rate. In addition, we consider out-of-pocket expenditures inflation. The rationale for including the out-of-pocket expenditures inflation is that prices of out-of-pocket purchases are easily observable in daily economic interactions. In line with the availability heuristic of Tversky and Kahnemann (1973), the easily recalled out-of-pocket expenditures inflation rate may thus give direction to how households perceive actual consumer price inflation. The table shows that the unemployment rate is weakly significant in the euro area previous to the introduction of the euro. It is highly significant for Sweden, both in quantified and quantitative survey data. While the money market rate is insignificant, out-of-pocket expenditures inflation is significant in the post-changeover period and in the quantified series for Sweden. The negative coefficient suggests that consumers over-react to out-of-pocket expenditures inflation in the sense that an increase in out-of-pocket inflation raises perceived inflation relative to actual inflation (which decreases the perception error).

Third, we investigate whether the bias in perceptions occurs because households are reluctant to revise their prior beliefs. This is the so called expectancy confirmation hypothesis investigated by Traut-Mattausch et al. (2004) in an experimental setting. As shown in the third panel of Table 2, past inflation expectations are significant only for the euro area and only in the post cash-changeover period.<sup>23</sup> The negative coefficient indicates that during this period, households' inflation perceptions overreacted to own past expectations such that perceptions exceeded actual inflation.<sup>24</sup>

The tests of information efficiency indicate that consumers could improve their inflation perceptions by using readily available information, such as past inflation or past perception errors. We conclude that inflation perceptions are not fully rational.

---

line with the findings for the EA aggregate.

<sup>23</sup>Household expectations of the inflation rate during the upcoming 12 months are also taken from the Joint Harmonized EU Consumer Survey. The qualitative response data is quantified using the probability method following Batchelor and Orr (1988). For Sweden, the mean of quantitative survey responses on expected inflation is available.

<sup>24</sup>This interpretation is confirmed by estimating a model that allows for different coefficients in periods of positive and negative perception errors. In both periods, the coefficient on expectations is negative.

Table 2: Information efficiency of inflation perceptions

	EA			SE	SEq
	93-07	93-01	03-07	93-07	96-07
<i>Serial correlation</i>					
$\pi_{t-12} - \pi_{t-12}^p$	0.3405** (0.1423)	0.5517*** (0.1442)	0.0199 (0.1687)	-0.3618*** (0.1372)	-0.4605*** (0.1215)
T	164	96	56	131	128
Adj. r-squared	0.47	0.45	0.47	0.13	0.23
<i>Strong-form efficiency</i>					
Unemployment rate (t-1)	0.0467 (0.0294)	0.0702* (0.0400)	-0.3212 (0.4308)	0.1053*** (0.0217)	0.0976** (0.0376)
Money market rate (t-1)	0.0214 (0.0435)	0.0651 (0.0446)	-0.2048 (0.3348)	-0.1033 (0.0648)	-0.0069 (0.0601)
Out-of-pocket $\pi_{t-1}$	-0.2820** (0.1094)	-0.0994 (0.0812)	-0.7435*** (0.1210)	-0.1209** (0.0496)	-0.0245 (0.0703)
$\pi_{t-1}$	0.9025*** (0.1652)	0.6380*** (0.1325)	1.9586*** (0.2600)	0.7214*** (0.0658)	0.5234*** (0.0718)
$\pi_{t-12} - \pi_t^{p-12}$	-0.1191 (0.0842)	-0.2627*** (0.0618)	-0.0717 (0.1370)	-0.2032** (0.0813)	-0.3119*** (0.0801)
T	127	59	56	127	127
Adj. r-squared	0.6	0.75	0.54	0.7	0.62
<i>Anchoring to expectations</i>					
$\pi_{t-12,t}^e$	-0.2042 (0.2682)	0.5584 (0.3422)	-1.2429*** (0.2185)	-0.0765 (0.3753)	0.0710 (0.3478)
$\pi_{t-12}^p$	-0.0015 (0.2671)	0.1727 (0.6966)	-0.9928*** (0.1349)	0.2364 (0.2222)	0.3023 (0.2871)
$\pi_{t-12}$	0.5083* (0.2756)	-0.1971 (0.3044)	0.4401 (0.2643)	-0.2895 (0.2761)	-0.4550*** (0.1206)
T	164	96	56	131	128
Adj. r-squared	0.51	0.53	0.65	0.17	0.23

*Notes:* This table investigates information efficiency of inflation perceptions in the euro area (EA) and Sweden (SE). Dependent variable is the perception error  $\pi_t - \pi_t^p$  based on quantified inflation perceptions for the EA and based on quantified (SE) and quantitative (SEq) inflation perceptions for Sweden. Estimations covering the 1993–2007 period allow for a level shift in 2002. OLS estimation with White standard errors in parentheses. \*, \*\* and \*\*\* indicate statistical significance at the 10%, 5% and 1% level, respectively.

### 4.3 Relation to Actual Inflation

Given that inflation perceptions are not fully rational, it is natural to ask whether and how perceptions are temporally related to actual inflation. If households update their judgements based on official HICP releases, then the HICP inflation rate should be Granger causal to perceived inflation. We expect a lagged effect of HICP figures since inflation numbers for a given month are published rather in the beginning to the middle of the following month, while consumers are surveyed already in the first three weeks of a month. Additionally, if consumers respond to inflation as observed in daily economic interactions, this should be reflected in an instantaneous relation between perceived and actual inflation.

We investigate Granger causality in the following bivariate vector-autoregression (VAR) with  $p + 1$  lags:

$$\begin{pmatrix} \pi_t^p \\ \pi_t \end{pmatrix} = y_t = A_0 + \sum_{i=1}^{p+1} A_i y_{t-i} + u_t$$

where  $A_0 = \begin{pmatrix} a_0^{11} \\ a_0^{21} \end{pmatrix}$ ,  $A_i = \begin{pmatrix} a_i^{11} & a_i^{12} \\ a_i^{21} & a_i^{22} \end{pmatrix}$  and  $u_t = \begin{pmatrix} u_{1,t} \\ u_{2,t} \end{pmatrix}$ . To test whether actual inflation is Granger-causal to perceived inflation we consider the Wald statistic that imposes the restriction  $a_1^{12} = \dots = a_p^{12} = 0$  on the first  $p$  lags in the estimated VAR( $p + 1$ ) model.<sup>25</sup> Instantaneous causality is assessed by testing whether contemporaneous residual correlation is zero.

The block-exogeneity tests reported in Table 3 suggest that actual inflation is Granger-causal to perceived inflation ( $\pi \rightarrow \pi^p$ ) in 6 out of 12 countries. A significant instantaneous relation ( $\pi \leftrightarrow \pi^p$ ) is detected in 8 countries. The table shows that a lagged or instantaneous relation exists in all countries except Italy and the Netherlands. For the euro area aggregate, both relations are highly significant. In sum, the results clearly indicate that households do not only react to reports on past inflation but also adjust their perceptions instantaneously to information that is available before the official HICP figures are released.

---

<sup>25</sup>It is well known that in the presence of highly persistent time series, the Wald statistic to assess Granger causality may follow a nonstandard distribution. This problem can be avoided by adding an extra lag that remains unrestricted when testing for causality, see Dolado and Lütkepohl (1996).

Table 3: Granger-causality in a bivariate VAR

<i>Country</i>	$\pi \leftrightarrow \pi^p$	$\pi \rightarrow \pi^p$	$\pi \leftarrow \pi^p$	Lags	T
AT	7.22 ***	1.41	0.61	2	141
BE	0.27	11.53 ***	1.36	2	174
DE	6.84 ***	2.07	0.01	2	138
EA	11.91 ***	12.97 ***	0.91	2	174
EL	3.05 *	0.42	0.02	2	174
ES	3.82 **	13.95 ***	0.31	2	174
FI	14.53 ***	3.24	1.45	3	139
FR	1.43	15.87 ***	0.00	2	173
IE	4.24 ***	6.15 **	0.47	2	138
IT	16.26 ***	6.15 **	5.65 **	2	174
NL	0.54	2.42	0.09	2	174
SE	14.29 ***	6.12 **	0.60	2	141
SEq	10.43 ***	2.31	0.01	2	138
UK	2.27	0.00	0.46	2	126

*Notes:* This table tests Granger-causality between perceived and actual inflation, 01/1993–07/2007. All estimations allow for a permanent level shift in 2002. To account for potential nonstationarity, the models are overfitted by including an extra lag not considered in block-exogeneity tests. *Lags* indicates the lag length as selected by Schwarz Bayesian information criterion (SBC). Instantaneous causality ( $\pi \leftrightarrow \pi^p$ ) is tested using the statistic  $T\hat{\rho}^2$  where  $\hat{\rho}$  is the contemporaneous correlation of residuals and  $T$  is the number of observations. Under the null hypothesis that  $\rho = 0$ , the statistic is asymptotically  $\chi^2(1)$  distributed. Granger causality ( $\pi \rightarrow \pi^p$ ,  $\pi \leftarrow \pi^p$ ) is assessed by testing for joint-exogeneity of lags of  $\pi$  and  $\pi^p$  in the equations for  $\pi^p$  and  $\pi$  respectively. The Wald statistic has a limiting  $\chi^2(lags)$ -distribution. \*, \*\* and \*\*\* indicate statistical significance at the 10%, 5% and 1% level, respectively.

#### 4.4 Implications for Belief Formation

The tests for information efficiency indicate that inflation perceptions are not rational. Also, we have documented that the survey mean of inflation perceptions is relatively inaccurate. Nevertheless, inflation perceptions respond to actual inflation, as suggested by the Granger-causality tests. For most countries, these tests suggest both a contemporaneous and a lagged response of perceptions to actual inflation. These patterns are broadly consistent with the epidemiological model of Carroll (2003). In this model, only a fraction of households update their beliefs in a given period. Consequently, the cross-sectional mean of inflation perceptions will fail rationality tests but is still contemporaneously related to actual inflation.

Further insights can be gained by considering the cross-sectional heterogeneity of inflation perceptions. As discussed in Section 3, Figure 1 shows the quantified cross-sectional standard deviation of perceptions in the euro area. The cross-sectional standard deviation averages at 1.29% which seems high given the moderate variability of inflation. As outlined

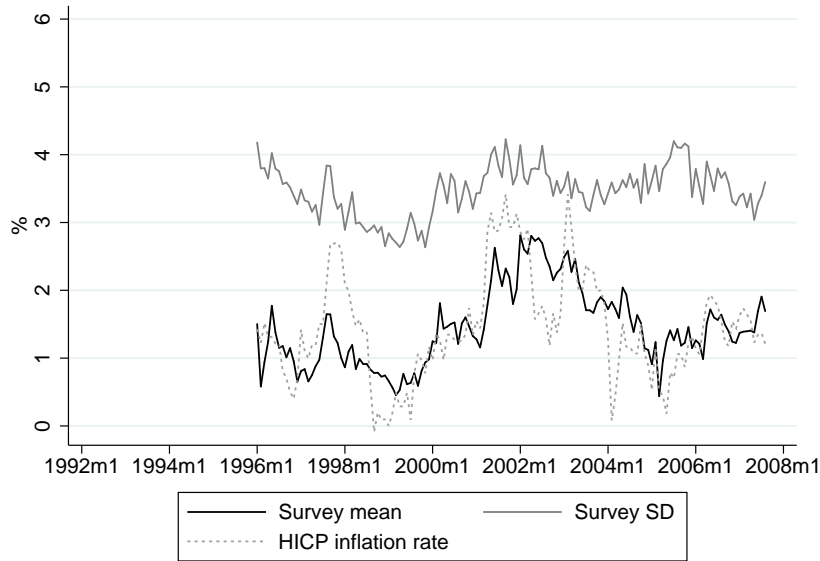


Figure 2: Mean and standard deviation of inflation perceptions in Sweden

*Notes:* This figure shows the mean and cross-sectional standard deviation of quantitative inflation perceptions from the Swedish Consumer Tendency Survey.

in Appendix A.2, the quantified standard deviation is even likely to underestimate the actual degree of heterogeneity. Figure 2 additionally shows the cross-sectional standard deviation of quantitative inflation perceptions taken from the Swedish Consumer Tendency Survey. Quantitative survey responses exhibit an average cross-sectional standard deviation of 3.46% during 1996–2007, whereas inflation perceptions average at 1.47% during the same period.

We thus arrive at a similar conclusion as Mankiw, Reis, and Wolfers (2004) who investigate inflation expectations of U.S. households. Inflation perceptions are not rational, yet related to contemporaneous and lagged actual inflation. The epidemiological model outlined in Section 2 is consistent with these broad patterns. Moreover, the model predicts a high degree of heterogeneity in perceptions, as the staggered updating mechanism generates heterogeneous information sets. The data confirms that inflation perceptions are highly heterogeneous. The next section thus investigates the epidemiological model in more detail.

## 5 Estimation of Epidemiological Models

### 5.1 Linear Partial Adjustment Models

This section discusses estimation results of the epidemiological models proposed in Section 2. Model (1) assumes that households update with contemporaneous actual inflation. The estimation equation is given by:

$$\pi_t^p = \alpha_0 + \alpha_1 \pi_t + \alpha_2 \pi_{t-1}^p + \varepsilon_t$$

Model (2) assumes that consumers who update refer to the most recent available official inflation figure, which is the one month lagged inflation rate. The regression equation reads:

$$\pi_t^p = \alpha_0 + \alpha_1 \pi_{t-1} + \alpha_2 \pi_{t-1}^p + \varepsilon_t$$

For the partial adjustment restriction to hold, we should not be able to reject the hypothesis that  $\alpha_1 + \alpha_2 = 1$ . We assess this restriction using a standard Wald test. To begin with, we estimate the models using ordinary least squares (OLS), employing White standard errors that allow for heteroskedasticity. The estimations for the sample period 1993–2007 control for the euro cash changeover by including an indicator variable that is unity during 2002–2007 and zero otherwise.

Tables 4 and 5 report country-by-country results for the period 1993–2007. Model (1) tends to fit marginally better than Model (2), as reflected in the R-squared and the high significance of the contemporaneous HICP inflation rate. The results for Model (2) suggest that in some countries not even a small proportion of consumers updates the perception of inflation with the lagged inflation rate. Both models are clearly rejected by the Wald test of the coefficient sum restriction. The respective p-values for the Wald tests of the coefficient restriction  $\alpha_1 + \alpha_2 = 1$  can be found in the second last column. If one nonetheless interprets the estimated coefficient on actual inflation as an updating frequency, estimations for Model (1) imply a quarterly updating frequency of 0.11 for the euro area.<sup>26</sup> This is considerably lower than the frequencies reported by studies that investigate inflation expectations. Using household expectations from the University of Michigan Survey of Consumers, Car-

---

<sup>26</sup>Quarterly frequencies are given by  $\lambda_{quarterly} = 1 - (1 - \lambda_{monthly})^3$ .

Table 4: Model (1), 1993–2007

<i>Country</i>	$\alpha_1$	$\alpha_2$	$\alpha_0$	$R^2$	Wald p	BG p
AT	0.0593*** (0.0217)	0.9120*** (0.0400)	0.0180 (0.0391)	0.97	0.41	0.06
BE	0.0436** (0.0192)	0.8338*** (0.0446)	0.1923*** (0.0521)	0.92	0.00	0.00
DE	0.0312** (0.0145)	0.9700*** (0.0295)	0.0023 (0.0379)	0.96	0.97	0.11
EA	0.0379*** (0.0121)	0.9000*** (0.0272)	0.1024*** (0.0339)	0.98	0.00	0.07
EL	0.0169** (0.0078)	0.8435*** (0.0440)	0.5731*** (0.1688)	0.92	0.00	0.00
ES	0.0632*** (0.0144)	0.8104*** (0.0328)	0.3116*** (0.0752)	0.96	0.00	0.73
FI	0.0483*** (0.0165)	0.8811*** (0.0384)	0.0743 (0.0543)	0.96	0.06	0.05
FR	0.0406*** (0.0139)	0.8637*** (0.0398)	0.1207*** (0.0438)	0.97	0.01	0.52
IE	0.0752*** (0.0177)	0.8268*** (0.0400)	0.2528*** (0.0715)	0.94	0.00	0.00
IT	0.0405** (0.0157)	0.8982*** (0.0317)	0.1267** (0.0556)	0.94	0.01	0.49
NL	0.0711*** (0.0206)	0.8482*** (0.0581)	0.1218 (0.0915)	0.90	0.12	0.00
SE	0.1314*** (0.0265)	0.6903*** (0.0477)	0.2318*** (0.0539)	0.81	0.00	0.68
SEq	0.1143*** (0.0303)	0.7657*** (0.0476)	0.1987*** (0.0607)	0.81	0.00	0.54
UK	0.0277 (0.0193)	0.8760*** (0.0425)	0.1519** (0.0610)	0.80	0.01	0.73

*Notes:* This table shows OLS estimates of Model (1). Sample periods are specified in Table A.1. The column *Wald p* reports the p-value of the Wald test of the restriction  $\alpha_1 + \alpha_2 = 1$ . *BG p* is the p-value of the Breusch-Godfrey LM test statistic for first order residual correlation. White standard errors allowing for heteroskedasticity in parentheses. \*, \*\* and \*\*\* indicate statistical significance at the 10%, 5% and 1% level, respectively.

Table 5: Model (2), 1993–2007

<i>Country</i>	$\alpha_1$	$\alpha_2$	$\alpha_0$	$R^2$	Wald p	BG p
AT	0.0356* (0.0210)	0.9222*** (0.0419)	0.0420 (0.0387)	0.97	0.22	0.04
BE	0.0493*** (0.0171)	0.8200*** (0.0458)	0.2045*** (0.0548)	0.92	0.00	0.01
DE	0.0177 (0.0130)	0.9663*** (0.0290)	0.0225 (0.0387)	0.96	0.62	0.09
EA	0.0224* (0.0126)	0.9217*** (0.0298)	0.0948*** (0.0358)	0.98	0.01	0.11
EL	0.0139* (0.0082)	0.8463*** (0.0450)	0.5779*** (0.1695)	0.92	0.00	0.00
ES	0.0546*** (0.0159)	0.8181*** (0.0363)	0.3176*** (0.0774)	0.96	0.00	0.88
FI	0.0223 (0.0170)	0.9014*** (0.0403)	0.0983* (0.0555)	0.96	0.05	0.02
FR	0.0359*** (0.0132)	0.8645*** (0.0382)	0.1268*** (0.0434)	0.97	0.00	0.32
IE	0.0752*** (0.0162)	0.8178*** (0.0383)	0.2829*** (0.0705)	0.94	0.00	0.01
IT	0.0246 (0.0161)	0.9176*** (0.0322)	0.1266** (0.0563)	0.94	0.01	0.47
NL	0.0688*** (0.0212)	0.8475*** (0.0578)	0.1303 (0.0925)	0.90	0.11	0.00
SE	0.0823*** (0.0276)	0.7421*** (0.0546)	0.2305*** (0.0589)	0.79	0.00	0.90
SEq	0.0732** (0.0323)	0.8029*** (0.0509)	0.2011*** (0.0623)	0.80	0.00	0.29
UK	0.0167 (0.0204)	0.8866*** (0.0453)	0.1497** (0.0630)	0.80	0.01	0.90

*Notes:* This table shows OLS estimates of Model (2). The sample periods are specified in Table A.1. The column *Wald p* reports the p-value of the Wald test of the restriction  $\alpha_1 + \alpha_2 = 1$ . *BG p* is the p-value of the Breusch-Godfrey LM test statistic for first order residual correlation. Estimates of the indicator variable for the euro cash changeover are not reported. White standard errors allowing for heteroskedasticity in parentheses. \*, \*\* and \*\*\* indicate statistical significance at the 10%, 5% and 1% level, respectively.



roll (2003) documents an updating frequency of 0.27. Döpke, Dovern, Fritsche, and Slacalek (2008b) estimate the updating parameter for four European countries and find updating frequencies that range between 0.18 and 0.32. In contrast to our results, Carroll (2003) and Döpke, Dovern, Fritsche, and Slacalek (2008b) report that the sum restriction cannot be rejected.<sup>27</sup> Tables A.10 through A.13 the Appendix additionally show results for the pre- and post-euro cash changeover periods. Levels as well as significances of estimated updating coefficients are mostly higher in the pre-euro cash changeover period. In all cases, the partial adjustment models are clearly rejected by the data.

Tables 4 and 5 also report the p-values of the Breusch-Godfrey LM test for first order serial correlation of residuals. For Model (1), the test signals significant residual correlation in 6 countries. For Model (2), significant residual correlation is detected in 7 countries. Due to the inclusion of a lagged dependent variable, residual correlation renders OLS inconsistent given that the lagged dependent variable is not predetermined anymore.<sup>28</sup> We thus additionally estimate a consistent specification that assumes an AR(1) residual process by the iterative Cochrane-Orcutt procedure. This procedure adjusts the original regression equation for first order serial correlation, such that the residuals of the resulting model are serially uncorrelated.<sup>29</sup> Tables A.14 and A.15 in the Appendix show the Cochrane-Orcutt estimation results for Models (1) and (2), respectively. The results are in line with the OLS estimation results. The parameter values have similar magnitudes and the Wald test generally rejects the hypothesis that  $\alpha_1 + \alpha_2 = 1$ .

All tables also report estimation results using the mean of quantitative survey perceptions from the Swedish Consumer Tendency Survey (denoted by the country code *SEq*). The

<sup>27</sup>More precisely, Döpke, Dovern, Fritsche, and Slacalek (2008b) report that the restriction is rejected for France, but holds for the other three countries as well as for the pooled sample.

<sup>28</sup>Given the model  $y_t = \beta_0 + \beta_1 y_{t-1} + u_t$ , OLS is only inconsistent if the condition  $E(y_{t-1} u_t) = 0$  is violated, i.e. if regressors are not predetermined. Note that theoretically, this condition can be satisfied even if residuals  $u_t = y_t - \beta_0 - \beta_1 y_{t-1}$  are serially correlated. This can be the case if  $u_t$  and  $y_{t-2}$  are correlated and  $E(y_{t-1} u_t) = 0$ . Then,  $E(u_t u_{t-1}) = E(u_t (y_{t-1} - \beta_0 - \beta_1 y_{t-2})) = -\beta_1 E(u_t y_{t-2}) \neq 0$ . In other words, consistency requires the special case that  $E(u_t u_{t-1}) = -\beta_1 E(u_t y_{t-2})$ .

<sup>29</sup>As outlined in Hamilton (1994), the Cochrane-Orcutt procedure for Model (1) converges to a (local) maximum of the following conditional likelihood function:

$$\begin{aligned} \mathcal{L} = & - \frac{T-1}{2} \log(2\pi) - \frac{T-1}{2} \log(\sigma^2) \\ & - \frac{1}{2\sigma^2} \sum_{t=2}^T (\pi_t^p - \alpha_0 - \alpha_1 \pi_t - \alpha_2 \pi_{t-1}^p - \rho [\pi_{t-1}^p - \alpha_0 - \alpha_1 \pi_{t-1} - \alpha_2 \pi_{t-2}^p])^2 \end{aligned}$$

The likelihood function for Model (2) is obtained by replacing  $\pi_t$  with  $\pi_{t-1}$ .

findings are consistent with results based on the quantified inflation perceptions, which corroborates the quantification method. In Sweden, the contemporaneous HICP inflation rate is highly relevant for perceived inflation with coefficient estimates of around 0.12. Compared to euro area countries, the results are relatively stable across subperiods.

Taking into account that some of the series are highly persistent, we also estimate the models in first differences. The estimation equation for Model (1) in first differences reads:

$$\Delta\pi_t^p = \alpha_0 + \alpha_1\Delta\pi_t + \alpha_2\Delta\pi_{t-1}^p + \varepsilon_t$$

Tables A.16 and A.17 in the Appendix report estimation results of Models (1) and (2), respectively. Both models are rejected even more clearly. The coefficients on actual inflation are of similar magnitude or slightly higher than in the estimations in levels. The coefficients on lagged inflation perceptions are mostly negative. Again, we obtain consistent results using the quantitative and qualitative response data on inflation perceptions from the Swedish survey.

As previously mentioned, it might well be the case that households do not refer to official HICP inflation when answering the survey. Rather they might report perceptions that rely on observed price changes in frequently bought items. We have estimated the models using out-of-pocket expenditures inflation as the measure of actual inflation. Results are qualitatively unchanged. Again the partial adjustment model is robustly rejected.<sup>30</sup>

We conclude that the two epidemiological models of perception formation are not adequate in our sample of European countries. A possible reason for this negative finding might be the assumption that the fraction of updating households is time-invariant. In the next section, we assess specifications that allow for time-varying adjustment parameters.

## 5.2 Non-Linear Adjustment

If households probabilistically update their information sets, then the probability of updating will likely be time-varying. In particular, one might expect that the probability of updating is higher if the inflation rate is high and if using outdated information becomes costly. This view is supported by empirical findings of Branch (2007). Branch (2007) shows that a model in which households rationally select predictors by optimizing costs and benefits of predicting inflation is consistent with inflation expectations from the University of Michigan Survey of

---

<sup>30</sup>Results are available upon request.

Table 6: Near-rationality in Model (1), 1993–2007

<i>Country</i>	$\alpha_1$	$\alpha_2$	$\Delta\alpha_1$	$\Delta\alpha_2$	T	Median				
AT	0.0046	(0.0404)	0.8761***	(0.0334)	-0.0023	(0.0159)	0.1081	(0.0751)	142	1.76
BE	-0.0103	(0.0461)	0.7996***	(0.0451)	0.0204	(0.0179)	0.0919*	(0.0542)	175	1.88
DE	0.0261	(0.0284)	0.9613***	(0.0433)	0.0120	(0.0139)	0.0038	(0.0450)	139	1.38
EA	0.0539**	(0.0255)	0.9020***	(0.0335)	-0.0075	(0.0085)	0.0011	(0.0347)	175	2.12
EL	-0.0083	(0.0844)	0.8078***	(0.0501)	-0.0501***	(0.0155)	0.0208	(0.0840)	175	3.81
ES	0.0770**	(0.0333)	0.8056***	(0.0352)	-0.0100	(0.0071)	-0.0739	(0.0508)	175	3.31
FI	0.0905*	(0.0505)	0.8861***	(0.0308)	-0.0158	(0.0263)	-0.0529	(0.0559)	141	1.36
FR	0.0372	(0.0235)	0.8437***	(0.0520)	0.0067	(0.0134)	-0.1233**	(0.0507)	175	1.77
IE	0.0333	(0.0330)	0.8485***	(0.0351)	-0.0206	(0.0200)	0.0682*	(0.0354)	140	2.65
IT	0.0032	(0.0468)	0.8806***	(0.0345)	-0.0102	(0.0129)	0.0408	(0.0497)	175	2.41
NL	0.0260	(0.0481)	0.8108***	(0.0596)	0.0023	(0.0176)	0.0353	(0.0500)	175	1.82
SE	0.1339***	(0.0482)	0.7085***	(0.0628)	-0.0434	(0.0316)	0.0231	(0.0603)	142	1.54
SEq	0.0977**	(0.0484)	0.7863***	(0.0613)	-0.0183	(0.0306)	0.0075	(0.0667)	139	1.54
UK	-0.0331	(0.0589)	0.8462***	(0.0489)	0.0006	(0.0172)	0.1103	(0.0688)	128	1.51

*Notes:* Parameters  $\Delta\alpha_1$  and  $\Delta\alpha_2$  denote the relative change in regression parameters in periods with actual inflation above median inflation. *Median* is the sample median of HICP inflation. Estimations allow for a level shift in 2002 and include separate constants for both regimes (not reported). White standard errors allowing for heteroskedasticity in parentheses. \*, \*\* and \*\*\* indicate statistical significance at the 10%, 5% and 1% level, respectively.

Consumers. The idea that economic agents only care about inflation if it becomes costly is formalized by Akerlof, Dickens, and Perry (2000). These authors introduce the concept of near rationality. In their model, near rational firms only fully incorporate expected inflation in wage and price-setting if ignoring inflation is sufficiently costly.

We thus allow for nonlinear updating in models with state-dependent adjustment parameters. We consider a simple specification that allows for different updating coefficients in periods of high and low inflation. Periods of high (low) inflation are characterized by actual HICP inflation which is above (below) the sample median of HICP inflation. As the median varies between countries, the estimations take into account that consumers in some countries are accustomed to higher median inflation rates than consumers in other countries. Tables 6 and A.18 summarize the estimation results for Models (1) and (2). The estimates of the interaction terms show that the coefficients on contemporaneous and lagged HICP inflation are generally not higher in periods of high inflation. Also in line with estimation results from the previous section, households in countries with relatively high median inflation rates (such as Ireland, Italy, Greece, Spain) do not show higher probabilities of updating. Hence, we find no evidence for non-linear adjustment or near-rationality in our sample of countries.<sup>31</sup>

### 5.3 Actual and Simulated Heterogeneity

The epidemiological model has direct implications for the cross-sectional heterogeneity of inflation perceptions. Hence, the model can also be tested by assessing the heterogeneity of inflation perceptions rather than the central tendency. For inflation expectations, Mankiw, Reis, and Wolfers (2004) show that the sticky information model is consistent with observed heterogeneity. Building on Mankiw, Reis, and Wolfers (2004), we compare the simulated cross-sectional heterogeneity of inflation perceptions in a population that is characterized by Model (1) to the actual heterogeneity of quantitative answers in the Swedish Consumer Tendency Survey. We measure heterogeneity by the cross-sectional standard deviation of quantitative survey responses. The updating parameter  $\lambda$  is set to 0.12, which corresponds to the coefficient estimate in the period 1996–2007. Figure 3 shows the simulated and survey based series. The mean perception of the simulated population is much smoother than actual mean of survey perceptions. More important, the cross-sectional standard deviation

---

<sup>31</sup>We have also tested for more sophisticated forms of non-linearity using the smooth transition regression framework. However, using the methods proposed by Teräsvirta (2004), linearity was generally not rejected.

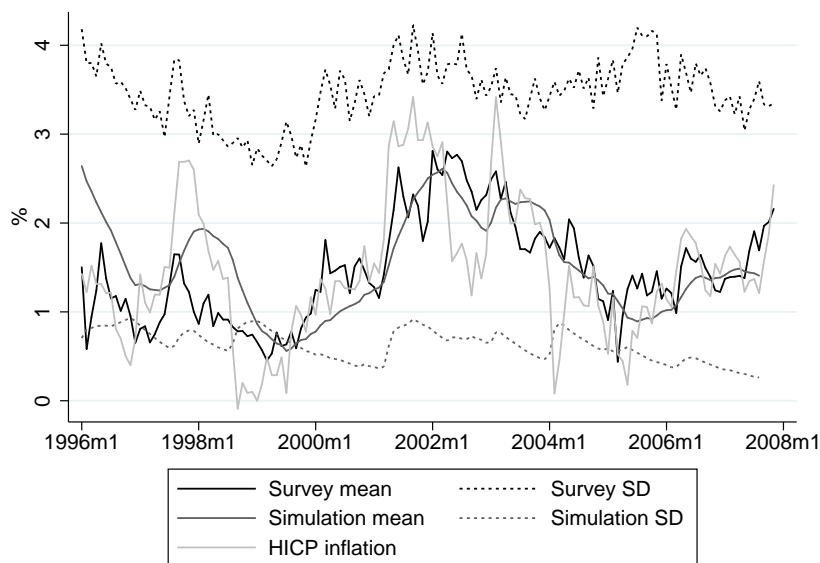


Figure 3: Model prediction and actual perceptions in Sweden

*Notes:* This figure shows actual and simulated mean and cross-sectional standard deviation (SD) of inflation perceptions from the Swedish Consumer Tendency Survey. The simulation is based on Model (1), i.e. households update with contemporaneous HICP inflation.  $\lambda = 0.12$ . The model is initialized in 01/1993.

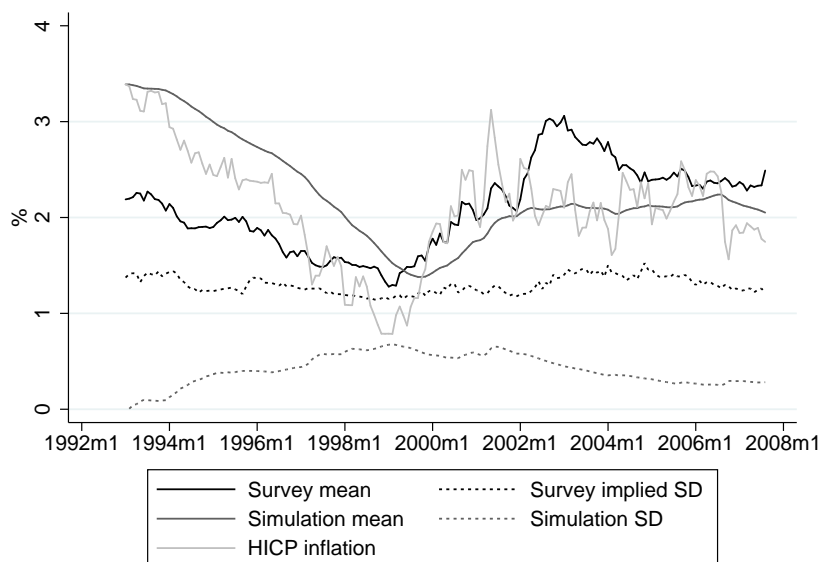


Figure 4: Model prediction and actual perceptions in the euro area

*Notes:* This figure shows the quantified and simulated mean and standard deviation (SD) of inflation perceptions in the euro area. The simulation is based on Model (1), i.e. households update with contemporaneous HICP inflation.  $\lambda = 0.05$ . The model is initialized in 01/1993.

of inflation perceptions in the simulated population is considerably lower than the standard deviation of actual quantitative survey responses. Also, the actual standard deviation does not show the distinct dynamic pattern induced by the epidemiological model. According to the model, heterogeneity rises following a persistent drop or surge in actual inflation to gradually decline again, as more and more individuals adjust their beliefs to the new level of inflation. This pattern is reflected in the simulation standard deviation but it is not visible in the actual survey standard deviation.

Consistent results are obtained for the euro area. Figure 4 shows the quantified standard deviation and the simulation standard deviation for the euro area aggregate. The updating parameter  $\lambda$  is set to 0.05. Again, the quantified standard deviation, which is likely to underestimate actual standard deviation, is much higher than the standard deviation of perceptions in the simulated population. Also, actual and simulated heterogeneity do not show common dynamics.

We conclude that the epidemiological model cannot explain the level and dynamics of cross-sectional heterogeneity. In particular, the level difference in heterogeneity suggests that other sources of heterogeneity exist than only infrequent updating. Potential sources of additional heterogeneity include that people update using different information or that people process information differently.

## 6 Conclusion

This paper investigates the dynamics of inflation perceptions in Europe. We use monthly household survey data from the Joint Harmonized EU Consumer Survey spanning 01/1993–08/2007. In an explorative investigation, we highlight three stylized facts about inflation perceptions. First, inflation perceptions do not efficiently incorporate available information and therefore fail rationality tests. Second, inflation perceptions are highly heterogeneous across the survey population. Third, inflation perceptions react both instantaneously and with a time lag to the actual rate of inflation.

These findings are broadly consistent with the epidemiological model of belief formation proposed by Carroll (2003). In this model, only a fraction of households update their information sets in a given month. The remaining households stay inattentive and stick to outdated beliefs about inflation. We estimate two epidemiological models of perception

formation, assuming that households either use the contemporaneous or the lagged rate of inflation to update their beliefs with. In almost all countries within the sample, however, both epidemiological models are clearly rejected as the sum of partial adjustment coefficients is significantly different from unity. This finding is in contrast to Carroll (2003) and Döpke, Dovern, Fritsche, and Slacalek (2008b). Using expectations data rather than perceptions data, the epidemiological model is not rejected by these authors.

The general picture is confirmed by quantitative response data from the Swedish Consumer Tendency Survey. Compared to euro area countries, the results are relatively stable across subperiods. Moreover, we find that estimations based on quantified qualitative survey data are consistent with estimations based on quantitative survey data. This corroborates the quantification method used for quantifying the qualitative response data from the Joint Harmonized EU Consumer Survey.

We conclude by comparing the cross-sectional heterogeneity of inflation perceptions in survey data with the heterogeneity of inflation perceptions in an artificial population that behaves according to an epidemiological model of belief formation. We find that the epidemiological model cannot explain the level and dynamics of survey heterogeneity. In particular, the model significantly underpredicts the actual level of heterogeneity. This suggests that other sources of heterogeneity exist than only infrequent updating. Potential sources of additional heterogeneity include that people update using different information or that people process information differently.

Looking ahead, further insights will be gained by investigating inflation perceptions across demographic groups. This will also contribute to the understanding of cross-sectional heterogeneity of beliefs. Moreover, for a better understanding of perception formation and to derive sensible models of aggregate dynamics, investigating household-level survey data will be essential.

## References

- AALTO-SETÄLÄ, V. (2006): “Why Do Consumers Believe That the Euro Raised Prices?,” *International Journal of Consumer Studies*, 30(2), 202–206.
- AKERLOF, G. A., W. T. DICKENS, AND G. L. PERRY (2000): “Near-Rational Wage and Price Setting and the Long-Run Phillips Curve,” *Brookings Papers on Economic Activity*, 31(1), 1–60.
- ALTISSIMO, F., M. EHRMANN, AND F. SMETS (2006): “Inflation Persistence and Price-Setting Behaviour in the Euro Area,” ECB Occasional Paper Series 46, European Central Bank.
- ANDOLFATTO, D., S. HENDRY, AND K. MORAN (2008): “Are Inflation Expectations Rational?,” *Journal of Monetary Economics*, 55(2), 406–422.
- ANG, A., G. BEKAERT, AND M. WEI (2007): “Do Macro Variables, Asset Markets, or Surveys Forecast Inflation Better?,” *Journal of Monetary Economics*, 54, 1163–1212.
- AUCREMANNE, L., M. COLLIN, AND T. STRAGIER (2007): “Assessing the Gap Between Observed and Perceived Inflation in the Euro Area: Is the Credibility of the HICP at Stake?,” National Bank of Belgium Working Paper 112.
- BANERJEE, A., M. MARCELLINO, AND C. OSBAT (2004): “Some Cautions on the Use of Panel Methods for Integrated Series of Macroeconomic Data,” *Econometrics Journal*, 7(2), 322–340.
- BATCHELOR, R. A., AND A. B. ORR (1988): “Inflation Expectations Revisited,” *Economica*, 55(219), 317–331.
- BENFORD, J., AND R. DRIVER (2008): “Public Attitudes to Inflation and Interest Rates,” *Bank of England Quarterly Bulletin*, 48(2), 148–156.
- BERK, J. M. (1999): “Measuring Inflation Expectations: A Survey Data Approach,” *Applied Economics*, 31(11), 1467–1480.
- BRANCH, W. A. (2007): “Sticky Information and Model Uncertainty in Survey Data on Inflation Expectations,” *Journal of Economic Dynamics and Control*, 31, 245–276.



- BREITUNG, J., AND M. H. PESARAN (2005): “Unit Roots and Cointegration in Panels,” Deutsche Bundesbank Discussion Paper 42/2005.
- BRYAN, M. F., AND G. VENKATU (2001a): “The Curiously Different Inflation Perspectives of Men and Women,” Federal Reserve Bank of Cleveland Economic Commentary Series.
- (2001b): “The Demographics of Inflation Opinion Surveys,” Federal Reserve Bank of Cleveland Economic Commentary Series.
- CARLSON, J. A., AND M. PARKIN (1975): “Inflation Expectations,” *Economica*, 42(166), 123–138.
- CARROLL, C. D. (2003): “Macroeconomic Expectations of Households and Professional Forecasters,” *Quarterly Journal of Economics*, 118(1), 269–298.
- DICKEY, D. A., AND W. A. FULLER (1981): “Likelihood Ratio Statistics for Autoregressive Time Series With a Unit Root,” *Econometrica*, 49(4), 1057–1072.
- DOEHRING, B., AND A. MORDONU (2007): “What Drives Inflation Perceptions? A Dynamic Panel Data Analysis,” European Commission Economic Papers 284.
- DOLADO, J., AND H. LÜTKEPOHL (1996): “Making Wald Tests Work for Cointegrated VAR Systems,” *Econometric Reviews*, 15, 369–386.
- DÖPKE, J., J. DOVERN, U. FRITSCHKE, AND J. SLACALEK (2008a): “The Dynamics of European Inflation Expectations,” *B.E. Journal of Macroeconomics*, 8(1).
- (2008b): “Sticky Information Phillips Curves: European Evidence,” *Journal of Money, Credit and Banking*, 40(7), 1513–1520.
- DZIUDA, W., AND G. MASTROBUONI (2006): “The Euro Changeover and Its Effects on Price Transparency and Inflation,” Working Paper 26, Collegio Carlo Alberto.
- ECB (2005): “Consumers’ Inflation Perceptions: Still At Odds With Official Statistics?,” *ECB Monthly Bulletin*, (4), 30–32.
- EHRBECK, T., AND R. WALDMANN (1996): “Why Are Professional Forecasters Biased? Agency Versus Behavioral Explanations,” *Quarterly Journal of Economics*, 111(1), 21–40.

- EHRMANN, M. (2006): “Rational Inattention, Inflation Developments and Perceptions After the Euro Cash Changeover,” ECB Working Paper 588, European Central Bank.
- EUROPEAN COMMISSION (2007): *The Joint Harmonised EU Programme of Business and Consumer Surveys: User Guide*, European Commission, Directorate-General for Economic and Financial Affairs.
- FLUCH, M., AND H. STIX (2007): “The Development of Euro Prices - Subjective Perception and Empirical Facts,” *Monetary Policy and the Economy (OeNB)*, 07(1), 55–84.
- FORSELLS, M., AND G. KENNY (2004): “Survey Expectations, Rationality and the Dynamics of Euro Area Inflation,” *Journal of Business Cycle Measurement and Analysis*, 1(1), 13–41.
- GfK (2002): *Consumer Tendency Survey*, GfK Sweden, Stockholm, Sweden.
- HAMILTON, J. D. (1994): *Time Series Analysis*. Princeton University Press.
- IM, K. S., M. H. PESARAN, AND Y. SHIN (2003): “Testing for Unit Roots in Heterogeneous Panels,” *Journal of Econometrics*, 115, 53–74.
- JONUNG, L. (1981): “Perceived and Expected Rates of Inflation in Sweden,” *American Economic Review*, 71(5), 961–968.
- JONUNG, L., AND D. LAIDLER (1988): “Are Perceptions of Inflation Rational? Some Evidence for Sweden,” *American Economic Review*, 78(5), 1080–1087.
- KHAN, H., AND Z. ZHU (2006): “Estimates of the Sticky-Information Phillips Curve for the United States,” *Journal of Money, Credit and Banking*, 38(1), 195–207.
- KWIATKOWSKI, D., P. C. B. PHILLIPS, P. SCHMIDT, AND Y. SHIN (1992): “Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root: How Sure Are We That Economic Time Series Have a Unit Root?,” *Journal of Econometrics*, 54(1-3), 159–178.
- LANNE, M., H. LÜTKEPOHL, AND P. SAIKKONEN (2002): “Comparison of Unit Root Tests for Time Series With Level Shifts,” *Journal of Time Series Analysis*, 23(6), 667–685.
- LASTER, D., P. BENNETT, AND I. S. GEOUM (1999): “Rational Bias in Macroeconomic Forecasts,” *Quarterly Journal of Economics*, 114(1), 293–318.

- LÜTKEPOHL, H., AND M. KRÄTZIG (eds.) (2004): *Applied Time Series Econometrics*. Cambridge University Press.
- MAAG, T. (2009): “On the Accuracy of the Probability Method for Quantifying Beliefs About Inflation,” KOF Working Papers 230, ETH Zurich.
- MACKINNON, J. G. (1991): “Critical Values for Cointegration Tests,” in *Long-Run Economic Relationships: Readings in Cointegration*, ed. by R. F. Engle, and C. W. J. Granger, pp. 267–276. Oxford University Press.
- MANKIW, N. G., AND R. REIS (2002): “Sticky Information Versus Sticky Prices: A Proposal to Replace the New Keynesian Phillips Curve,” *Quarterly Journal of Economics*, 117(4), 1295–1328.
- (2006): “Pervasive Stickiness,” *American Economic Review*, 96(2), 164–169.
- MANKIW, N. G., R. REIS, AND J. WOLFERS (2004): “Disagreement About Inflation Expectations,” *NBER Macroeconomics Annual 2003*, 18, 209–248.
- MEHRA, Y. P. (2002): “Survey Measures of Expected Inflation: Revisiting the Issues of Predictive Content and Rationality,” *Economic Quarterly*, 88(3), 17–36.
- O’REILLY, G., AND K. WHELAN (2005): “Has Euro-Area Inflation Persistence Changed Over Time?,” *Review of Economics and Statistics*, 87(4), 709–720.
- PALMQVIST, S., AND L. STRÖMBERG (2004): “Households’ Inflation Opinions - A Tale of Two Surveys,” *Sveriges Riksbank Economic Review*, 4, 23–42.
- PEDRONI, P. (1999): “Critical Values for Cointegration Tests in Heterogeneous Panels With Multiple Regressors,” *Oxford Bulletin of Economics and Statistics*, 61, 653–670.
- (2004): “Panel Cointegration: Asymptotic and Finite Sample Properties of Pooled Time Series Tests With an Application to the PPP Hypothesis,” *Econometric Theory*, 20, 597–625.
- PERRON, P. (1988): “Trends and Random Walks in Macroeconomic Time Series,” *Journal of Economic Dynamics and Control*, 12, 297–332.

- PESARAN, M. H. (2007): “A Simple Panel Unit Root Test in the Presence of Cross-Section Dependence,” *Journal of Applied Econometrics*, 22(2), 265–312.
- PIERSE, R. G., AND A. J. SNELL (1995): “Temporal Aggregation and the Power of Tests for a Unit Root,” *Journal of Econometrics*, 65(2), 333–429.
- SAIKKONEN, P., AND H. LÜTKEPOHL (2002): “Testing for a Unit Root in a Time Series With a Level Shift at Unknown Time,” *Econometric Theory*, 18(2), 313–348.
- STIX, H. (2005): “Perceived Inflation and the Euro: Why High? Why Persistent?,” in *Proceedings of OeNB Workshops: Price Setting and Inflation Persistence, December 15, 2005*, no. 8, pp. 221–249. Oesterreichische Nationalbank.
- TERÄSVIRTA, T. (2004): “Smooth Transition Regression Modelling,” in *Applied Time Series Econometrics*, ed. by H. Lütkepohl, and M. Kräätzig, chap. 6, pp. 222–242. Cambridge University Press.
- THEIL, H. (1952): “On the Shape of Microvariables and the Munich Business Test,” *Revue de l’Institut International de Statistique*, 20, 105–120.
- THOMAS, L. B. (1999): “Survey Measures of Expected U.S. Inflation,” *Journal of Economic Perspectives*, 13(4), 125–144.
- TRAUT-MATTAUSCH, E., S. SCHULZ-HARDT, T. GREITEMEYER, AND D. FREY (2004): “Expectancy Confirmation in Spite of Disconfirming Evidence: The Case of Price Increases due to the Introduction of the Euro,” *European Journal of Social Psychology*, 34, 739–760.
- TVERSKY, A., AND D. KAHNEMAN (1973): “Availability: A Heuristic for Judging Frequency and Probability,” *Cognitive Psychology*, 5(2), 207–232.
- WANG, P., AND Y. WEN (2007): “Inflation Dynamics: A Cross-Country Investigation,” *Journal of Monetary Economics*, 54, 2004–2031.

# Appendix

## A.1 Sample Summary

Table A.1: Sample summary

Code	Country	Sample		T	Mean $\pi$	Median $\pi$	SD $\pi$
AT	Austria	10/1995	– 08/2007	143	1.58	1.68	0.60
BE	Belgium	01/1993	– 08/2007	176	1.87	1.88	0.71
DE	Germany	01/1993	– 08/2007	176	1.41	1.38	0.56
EA	Euro area	01/1993	– 08/2007	176	2.11	2.12	0.58
EL	Greece	01/1993	– 08/2007	176	5.33	3.81	3.30
ES	Spain	01/1993	– 08/2007	176	3.28	3.31	1.00
EU	Europe	01/1993	– 08/2007	176	3.21	2.95	1.37
FI	Finland	11/1995	– 08/2007	142	1.43	1.32	0.85
FR	France	01/1993	– 07/2007	175	1.71	1.77	0.59
IE	Ireland	01/1993	– 08/2007	176	3.00	2.65	1.25
IT	Italy	01/1993	– 08/2007	176	2.85	2.41	1.15
NL	Netherlands	01/1993	– 08/2007	176	2.14	1.82	1.07
SE	Sweden	10/1995	– 08/2007	143	1.48	1.37	0.79
UK	United Kingdom	01/1993	– 08/2007	176	1.57	1.51	0.54

*Notes:* The last three columns show mean, median and standard deviation of the HICP inflation rate in corresponding sample periods. The sample generally spans 01/1993 to 08/2007 and is defined by the joint availability of survey data and HICP inflation rates.  $T$  denotes the number of monthly observations.

## A.2 Quantifying Inflation Perceptions

The Joint Harmonized EU Consumer Survey captures perceived inflation by asking: “How do you think that consumer prices have developed over the last 12 months? They have...”. Answers are given on an ordinal scale: “Fallen ( $S_1$ ), stayed about the same ( $S_2$ ), risen slightly ( $S_3$ ), risen moderately ( $S_4$ ), risen a lot ( $S_5$ )”. For further reference,  $S_1$  through  $S_5$  denote the answer categories, whereas  $s_1$  through  $s_5$  are the share of responses in the corresponding category excluding the additional “don’t know”-category. We quantify the qualitative response data employing the 5-category probability method.

Assume that previous to answering the survey, respondent  $i$  forms a quantitative perception  $\pi_{it}^p$  of inflation. Respondent  $i$  then answers the EU consumer survey question according to the following scheme:

$$\begin{aligned}
\pi_{t,i}^p < -\delta_t & : \text{ prices in general are lower } (S_1) \\
-\delta_t \leq \pi_{t,i}^p < \delta_t & : \text{ about the same } (S_2) \\
\delta_t \leq \pi_{t,i}^p < \pi_t^r - \eta_t & : \text{ a little higher } (S_3) \\
\pi_t^r - \eta_t \leq \pi_{t,i}^p < \pi_t^r + \eta_t & : \text{ moderately higher } (S_4) \\
\pi_{t,i}^p \geq \pi_t^r + \eta_t & : \text{ a lot higher } (S_5)
\end{aligned}$$

Judgements  $\pi_{it}^p$  will vary across respondents due to differences in information processing or due to the use of different information sets. Assuming that  $\pi_{it}^p \sim N(\pi_t^p, \sigma_t)$  and that range parameters  $\delta_t$  and  $\eta_t$  are homogeneous across respondents we obtain:

$$\begin{aligned}
s_t^1 &= P(\pi_{t,i}^p < -\delta_t) = \Phi\left(\frac{-\delta_t - \pi_t^p}{\sigma_t}\right) \\
s_t^2 &= P(-\delta_t \leq \pi_{t,i}^p < \delta_t) = \Phi\left(\frac{\delta_t - \pi_t^p}{\sigma_t}\right) - \Phi\left(\frac{-\delta_t - \pi_t^p}{\sigma_t}\right) \\
s_t^3 &= P(\delta_t \leq \pi_{t,i}^p < \pi_t^r - \eta_t) = \Phi\left(\frac{\pi_t^r - \eta_t - \pi_t^p}{\sigma_t}\right) - \Phi\left(\frac{\delta_t - \pi_t^p}{\sigma_t}\right) \\
s_t^4 &= P(\pi_t^r - \eta_t \leq \pi_{t,i}^p < \pi_t^r + \eta_t) = \Phi\left(\frac{\pi_t^r + \eta_t - \pi_t^p}{\sigma_t}\right) - \Phi\left(\frac{\pi_t^r - \eta_t - \pi_t^p}{\sigma_t}\right) \\
s_t^5 &= P(\pi_t^r + \eta_t \leq \pi_{t,i}^p) = 1 - \Phi\left(\frac{\pi_t^r + \eta_t - \pi_t^p}{\sigma_t}\right)
\end{aligned} \tag{A.1}$$

where  $\Phi(\cdot)$  is the standard normal cumulative distribution function. The system of equations (A.1) can be rewritten to obtain a system of 4 linearly independent equations with 5 unknowns  $(\pi_t^p, \sigma_t, \delta_t, \eta_t, \pi_t^r)$ :

$$\begin{aligned}
G_t^1 &= \Phi^{-1}(s_t^1) = \frac{-\delta_t - \pi_t^p}{\sigma_t} \\
G_t^2 &= \Phi^{-1}(1 - s_t^5 - s_t^4 - s_t^3 - s_t^2) = \frac{-\delta_t - \pi_t^p}{\sigma_t} \\
G_t^3 &= \Phi^{-1}(1 - s_t^5 - s_t^4 - s_t^3) = \frac{\delta_t - \pi_t^p}{\sigma_t} \\
G_t^4 &= \Phi^{-1}(1 - s_t^5 - s_t^4) = \frac{\pi_t^r - \eta_t - \pi_t^p}{\sigma_t} \\
G_t^5 &= \Phi^{-1}(1 - s_t^5) = \frac{\pi_t^r + \eta_t - \pi_t^p}{\sigma_t}
\end{aligned} \tag{A.2}$$

System (A.2) can be solved for the mean  $\pi_t^p$  of perceived inflation:

$$\pi_t^p = \pi_t^r \frac{G_t^2 + G_t^3}{G_t^2 + G_t^3 - G_t^4 - G_t^5} \tag{A.3}$$

In the following,  $\pi_t^p$  is referred to as “perceived inflation” (rather than “mean of perceived inflation”). The remaining unknowns are given by:

$$\sigma_t = \pi_t^r \frac{-2}{G_t^2 + G_t^3 - G_t^4 - G_t^5} \tag{A.4}$$

$$\delta_t = \pi_t^r \frac{G_t^2 - G_t^3}{G_t^2 + G_t^3 - G_t^4 - G_t^5} \tag{A.5}$$

$$\eta_t = \pi_t^r \frac{G_t^4 - G_t^5}{G_t^2 + G_t^3 - G_t^4 - G_t^5} \tag{A.6}$$

To identify the above system, we assume that the reference rate of inflation  $\pi_t^r$  (the “moderate” rate of inflation) is constant over time but may differ across countries. Hence,  $\pi_t^r = \pi^r$  is a constant scaling factor to perceived inflation. To determine the moderate level of inflation, we impose unbiasedness of perceived inflation such that average perceived inflation is equal to average actual inflation over the sample period:<sup>32</sup>

$$\pi_t^r = \frac{\bar{\pi}}{\frac{1}{T} \sum_{t=1}^T \frac{G_t^2 + G_t^3}{G_t^2 + G_t^3 - G_t^4 - G_t^5}} \tag{A.7}$$

where  $T$  is the number of periods and  $\bar{\pi}$  the average actual rate of inflation.

---

<sup>32</sup>The unbiasedness assumption is commonly imposed to quantify inflation expectations, see, e.g., Berk (1999) and Forsells and Kenny (2004).

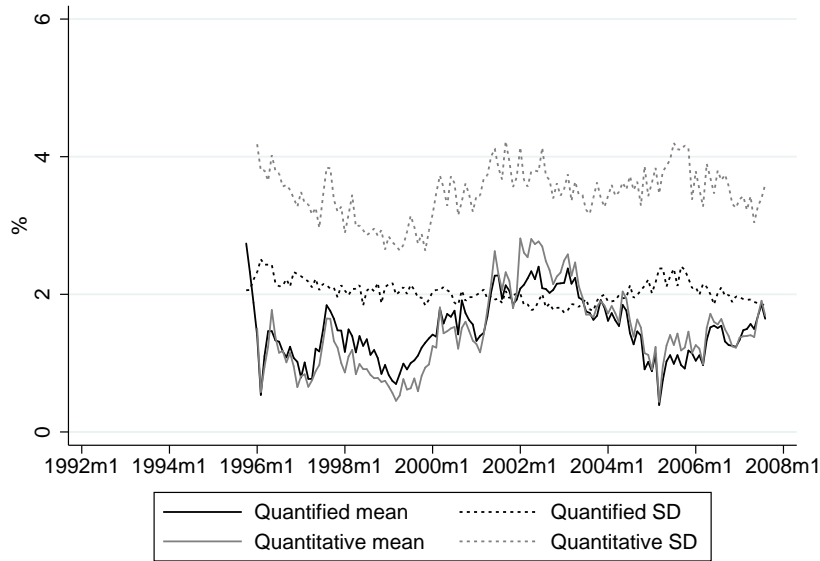


Figure A.1: Perceived inflation in Sweden

*Notes:* The figure shows quantified mean and cross-sectional standard deviation of inflation perceptions as well as the mean and standard deviation of quantitative inflation perceptions. Qualitative response data is quantified using the 5-category probability method under the assumption that perceptions are unbiased.

The assumptions imposed by the probability approach have been critically discussed in the literature. To assess the method, Figure A.1 shows quantified inflation perceptions as well as actual quantitative perceptions which are available from the Swedish Consumer Tendency Survey. The quantified mean closely tracks the mean of quantitative survey responses. The correlation coefficient of the two series is 0.96. The level difference averages at 0.01%. Quantitative response data is also available for Austria, where a survey was conducted in June 2004. Stix (2005) reports that inflation perceptions average at 2.7%. The probability method generates a value of 2.20%. Figure A.1 further indicates that the quantified cross-sectional standard deviation of perceptions is less accurate. The correlation of the quantified series with the standard deviation of quantitative responses is only 0.19. Moreover, the quantified standard deviation averages 1.43% below the actual standard deviation of quantitative responses. For assessing the heterogeneity generated by models of belief formation we therefore primarily rely on the quantitative response data from Sweden.



### A.3 Time Series Properties

Table A.2: Stationarity of the HICP inflation rate

<i>Country</i>	<i>T</i>	<i>p</i>	$\hat{\tau}_T$	$\Phi_3$	$\hat{\tau}_\mu$	$\Phi_1$	$\hat{\tau}$	Levels	1st. diff.	KPSS
AT	142	1	-2.52	3.19	-2.07	2.20	-0.45	I(1)	I(0)	***
BE	175	1	-3.29 *	5.40	-3.26 **	5.38 **	-1.36	I(0), c	I(0)	***
DE	175	1	-2.84	4.97	-3.08 **	4.89 *	-1.74 *	I(0), c	I(0)	***
EA	175	1	-2.20	2.47	-2.17	2.46	-1.09	I(1)	I(0)	***
EL	175	1	-1.91	3.61	-2.66 *	5.57 **	-3.14 ***	I(0), c	I(0)	***
ES	175	1	-2.47	3.05	-2.26	2.62	-1.00	I(1)	I(0)	***
EU	175	12	-1.94	2.57	-2.22	2.58	-1.00	I(1)	I(0)	***
FI	141	2	-1.79	1.66	-1.72	1.51	-0.61	I(1)	I(0)	***
FR	175	12	-2.24	2.52	-2.15	2.35	-0.86	I(1)	I(0)	**
IE	175	1	-2.21	2.46	-2.22	2.47	-0.79	I(1)	I(0)	***
IT	175	1	-1.85	1.76	-1.62	1.70	-1.40	I(1)	I(0)	***
NL	127	1	-1.72	1.90	-1.54	1.19	-0.68	I(1)	I(0)	***
SE	142	1	-3.50 **	6.48 ***	-3.60 ***	6.62 **	-2.12 ***	I(0), ct	I(0)	***
UK	175	1	-2.77	3.84	-2.77 *	3.86 *	-0.69	I(1)	I(0)	***

*Notes:* This table shows ADF and KPSS tests on country level. The sample periods are specified in Table A.1.  $T$  is the number of observations. Following Fuller (1976) and Dickey and Fuller (1981),  $\hat{\tau}_T$  denotes the t-statistic in the specification with constant and deterministic trend,  $\Phi_3$  is the F-statistic for the joint hypothesis that coefficient and time trend are zero.  $\hat{\tau}_\mu$  denotes the t-statistic in the specification with constant only,  $\Phi_1$  is the F-statistic for the joint test that coefficient and constant are zero.  $\hat{\tau}$  denotes the t-statistic in the specification without constant and trend. Columns *Levels* and *1st. diff.* show the model specification implied by the sequential procedure of Perron (1988). The number of lags  $p$  is determined using the Schwarz Bayesian information criterion (SBC). Critical values according to MacKinnon (1991) and Dickey and Fuller (1981). The KPSS test allows for deterministic trend, critical values are taken from Kwiatkowski et al. (1992). \*, \*\* and \*\*\* indicate statistical significance at the 10%, 5% and 1% level, respectively.

Table A.3: Stationarity of perceived inflation

Country	T	p	$\hat{\tau}_\tau$	$\Phi_3$	$\hat{\tau}_\mu$	$\Phi_1$	$\hat{\tau}$	Levels	1st. diff.	KPSS
AT	143	1	-2.05	2.26	-1.36	2.38	-0.95	I(1)	I(0)	0.81 ***
BE	176	1	-2.60	4.55	-1.11	0.66	-0.10	I(1)	I(0)	1.44 ***
DE	176	1	-1.93	2.59	-1.55	1.21	-0.33	I(1)	I(0)	0.77 ***
EA	176	1	-1.70	1.81	-0.91	0.46	-0.10	I(1)	I(0)	1.09 ***
EL	176	1	-2.17	3.48	-0.68	0.43	-0.44	I(1)	I(0)	1.54 ***
ES	176	1	-1.81	1.63	-1.29	0.89	-0.01	I(1)	I(0)	1.22 ***
EU	176	1	-1.84	2.14	-1.13	0.67	-0.07	I(1)	I(0)	1.17 ***
FI	142	1	-5.50 ***	19.93 ***	-5.94 ***	21.45 ***	-0.14	I(0), ct	I(0)	0.96 ***
FR	175	1	-2.17	2.36	-1.21	1.10	-0.44	I(1)	I(0)	0.97 ***
IE	176	1	-1.48	1.53	-1.73	1.57	-0.02	I(1)	I(0)	0.96 ***
IT	176	1	-2.49	3.23	-2.44 **	2.99	-0.25	I(1)	I(0)	0.83 ***
NL	176	1	-1.37	1.14	-1.51	1.14	-0.51	I(1)	I(0)	0.86 ***
SE	143	1	-3.28 **	5.72 *	-3.14 ***	4.99 **	-1.15	I(0), ct	I(0)	0.68 ***
UK	176	1	-2.49	3.23	-2.44	2.99	-0.25	I(1)	I(0)	1.03 ***

Notes: See footnote of Table A.2 for a description.

Table A.4: Unit root tests allowing for a deterministic level shift

<i>Country</i>	HICP inflation rate		Perceived inflation	
	Break date	Test statistic	Break date	Test statistic
AT	04/2001	-2.89 **	06/2000	-2.43
BE	08/2000	-2.08	06/2000	-1.71
DE	09/2006	-1.71	01/2002	-1.90
EA	01/2002	-1.78	02/2002	-2.16
EL	09/1993	-3.08 **	06/2005	-1.10
ES	07/2001	-1.68	04/2003	-0.94
EU	01/2002	-2.16	02/2002	-2.58 *
FI	01/1994	-2.27	08/1997	-4.27 ***
FR	01/2002	-2.27	09/1995	-1.00
IE	12/2000	-1.50	12/2004	-1.87
IT	01/2002	-1.81	11/2004	-1.63
NL	01/2001	-0.44	10/2002	-1.18
SE	01/1994	-2.69 **	03/2005	-2.18
UK	10/1995	-3.02 **	03/2006	-2.45

*Notes:* This table shows unit root test results for actual HICP inflation and perceived inflation. Following Saikkonen and Lütkepohl (2002) and Lanne, Lütkepohl, and Saikkonen (2002) the tests allow for an exponential level shift in the data generating process. The sample periods are specified in Table A.1.  $T$  is the number of observations. The test is based on estimating the deterministic term first by generalized least squares under the null hypothesis of a unit root. Subsequently, an ADF type test is performed on the adjusted series which also includes terms to correct for estimation errors in the parameters of the deterministic part. The exponential shift begins at the break date, which is chosen to minimize the generalized sum of squared residuals. Setting the break date exogenously to 01/2002 does not change any of the results. Critical values are taken from Lanne, Lütkepohl, and Saikkonen (2002). Estimation is done with the JMulti software from Lütkepohl and Krätzig (2004). \*, \*\* and \*\*\* indicate statistical significance at the 10%, 5% and 1% level, respectively.

Table A.5: Panel unit root tests

		p=1	p=2	p=3	p=4
<i>1993–2007</i>					
$\pi$ ( $N=12$ , $T=175$ )	IPS	-2.520*	-2.465*	-2.758**	-2.766**
	CIPS	-2.866**	-2.780**	-3.085***	-3.141***
$\pi^p$ ( $N=9$ , $T=175$ )	IPS	-2.192	-2.063	-2.063	-2.048
	CIPS	-2.409	-2.246	-2.206	-2.068
<i>1993–2001</i>					
$\pi$ ( $N=12$ , $T=108$ )	IPS	-2.222	-2.229	-2.501*	-2.567**
	CIPS	-2.473	-2.432	-2.721*	-2.978***
$\pi^p$ ( $N=9$ , $T=108$ )	IPS	-2.401	-2.211	-2.180	-2.211
	CIPS	-2.545	-2.364	-2.205	-2.123
<i>2003–2007</i>					
$\pi$ ( $N=12$ , $T=55$ )	IPS	-2.528*	-2.042	-2.070	-1.944
	CIPS	-2.527	-2.178	-2.098	-1.942
$\pi^p$ ( $N=12$ , $T=55$ )	IPS	-2.637**	-2.464*	-2.525**	-2.302
	CIPS	-2.803**	-2.661*	-2.405	-2.512

*Notes:* This table shows IPS and CIPS panel unit root tests for actual HICP inflation  $\pi$  and perceived inflation  $\pi^p$ . IPS denotes the Im, Pesaran and Shin (2003)  $t$ -bar statistic and accounts for common time effects. Following Pesaran (2007), CPIS is the  $t$ -bar statistic based on cross-sectionally augmented ADF regressions. Critical values are provided in the respective papers. All statistics are based on AR( $p$ ) specifications in levels that include a deterministic trend and a constant. To obtain balanced panels for perceived inflation, AT, FI, SE are excluded in the samples 1993–2007 and 1993–2001. \*, \*\* and \*\*\* indicate statistical significance at the 10%, 5% and 1% level, respectively.

Table A.6: Cointegration of actual and perceived inflation, 1993–2007

<i>Country</i>	r=0	r=1	Implied r	$\beta$	p	T		
AT	10.49	1.78	0	0.74	2	140		
BE	14.36	1.57	0	0.26	2	173		
DE	11.99	2.56	0	0.15	2	173		
EA	6.44	0.91	0	0.19	2	173		
EL	15.93	**	0.33	1	-0.06	3	172	
ES	6.75	1.62	0	0.25	2	173		
EU	9.89	1.40	0	0.25	2	173		
FI	39.28	***	4.05	**	2	0.21	2	139
FR	14.58	1.46	0	0.52	2	173		
IE	16.01	**	3.49	1	0.36	2	173	
IT	5.07	2.23	0	0.23	2	173		
NL	14.53	4.97	**	0	0.24	2	125	
SE	26.63	***	9.60	***	2	0.37	2	140
UK	15.32	4.40	**	0	0.13	2	173	

*Notes:* This table shows Johansen tests for the cointegration rank between actual and perceived inflation, 1993–2007. The lag order  $p$  is determined using the SBC, a minimum of one lag in first differences is included. Critical values from Johansen (1995):  $r=1$ , 15.41 (5%, \*\*), 19.96 (1%, \*\*\*);  $r=2$ , 3.76 (5%), 9.24 (1%).  $\beta$  is the cointegration coefficient in the regression  $y_{it} = \alpha + \beta x_{it} + \epsilon_{it}$ .  $T$  denotes the number of observations.

Table A.7: Cointegration of actual and perceived inflation, 1993–2001

<i>Country</i>	r=0		r=1		Implied r	$\beta$	p	T
AT	6.92		1.16		0	0.55	2	73
BE	20.01	***	3.27		1	0.28	2	106
DE	9.91		3.82		0	0.26	2	106
EA	5.70		1.81		0	0.25	2	106
EL	12.56		3.33		0	0.12	2	106
ES	10.28		2.41		0	0.31	2	106
EU	8.54		2.40		0	0.32	2	106
FI	33.30	***	3.30		1	0.60	2	72
FR	8.19		2.84		0	0.27	2	106
IE	9.31		1.49		0	0.39	2	106
IT	6.94		1.83		0	0.35	2	106
NL	12.99		0.11		0	0.27	2	58
SE	22.01	***	6.30	**	2	0.33	2	73
UK	8.39		0.82		0	0.10	2	106

*Notes:* See footnote of Table A.6 for a description.

Table A.8: Cointegration of actual and perceived inflation, 2003–2007

<i>Country</i>	r=0		r=1		Implied r	$\beta$	p	T
AT	10.74		4.48	**	0	0.05	2	53
BE	10.41		3.86	**	0	0.12	2	53
DE	27.04	***	7.56	***	2	-0.24	2	53
EA	16.75	**	6.42	**	2	-0.01	2	53
EL	23.58	***	9.50	***	2	-0.55	2	53
ES	11.49		5.68	**	0	0.09	2	53
EU	26.62	**	3.07		1	-0.23	2	53
FI	12.11		2.99		0	0.14	2	53
FR	18.09	**	6.36	**	2	0.07	2	53
IE	18.65	**	4.94	**	2	0.34	2	53
IT	10.37		0.64		0	1.51	2	53
NL	22.75	***	6.18	**	2	1.12	2	53
SE	20.83	***	7.20	***	2	0.48	2	53
UK	11.44		2.02		0	0.28	2	53

*Notes:* See footnote of Table A.6 for a description.

Table A.9: Tests for panel cointegration

	1993–2007	1993–2001	2003–2007
<i>Parametric t-statistic</i>			
Panel statistic	-1.62**	-3.59***	-2.73***
Group mean statistic	-2.94***	-4.31***	-2.98***
<i>Nonparametric t-statistic</i>			
Panel statistic	-3.09***	-5.28***	-2.98***
Group mean statistic	-4.52***	-6.18***	-4.36***

*Notes:* Pedroni tests for panel cointegration of actual inflation and perceived inflation. Panel cointegration regressions include time fixed effects. All statistics are standardized and follow a  $N(0, 1)$  distribution, see Pedroni (1999, 2004). Estimation has been done using the RATS procedure written by Peter Pedroni. \*, \*\* and \*\*\* indicate statistical significance at the 10%, 5% and 1% level, respectively.

## A.4 Further Results



Table A.10: Model (1), 1993–2001

Country	$\alpha_1$	$\alpha_2$	$\alpha_0$	$R^2$	Wald p	BG p
AT	0.0722** (0.0273)	0.9211*** (0.0447)	-0.0090 (0.0447)	0.93 (0.0406)	0.85	0.35
BE	0.0619** (0.0275)	0.7957*** (0.0571)	0.2221*** (0.0571)	0.85 (0.0576)	0.00	0.01
DE	0.0761*** (0.0258)	0.7997*** (0.0621)	0.1519*** (0.0621)	0.92 (0.0566)	0.01	0.90
E.A	0.0635*** (0.0156)	0.8301*** (0.0362)	0.1766*** (0.0362)	0.96 (0.0426)	0.00	0.04
EL	0.0235*** (0.0082)	0.8065*** (0.0467)	0.6954*** (0.0467)	0.85 (0.1753)	0.00	0.52
ES	0.0606*** (0.0168)	0.8186*** (0.0409)	0.2977*** (0.0409)	0.89 (0.0887)	0.00	0.44
FI	0.0364 (0.0248)	0.9005*** (0.0399)	0.0740 (0.0399)	0.96 (0.0615)	0.12	0.09
FR	0.0489*** (0.0179)	0.8813*** (0.0474)	0.0851* (0.0474)	0.86 (0.0482)	0.08	0.53
IE	0.0858*** (0.0232)	0.8050*** (0.0542)	0.2802*** (0.0542)	0.96 (0.0895)	0.00	0.29
IT	0.0468** (0.0208)	0.8573*** (0.0471)	0.2115*** (0.0471)	0.92 (0.0704)	0.00	0.87
NL	0.0959*** (0.0192)	0.6509*** (0.0624)	0.4213*** (0.0624)	0.88 (0.0832)	0.00	0.77
SE	0.1644*** (0.0320)	0.5396*** (0.0668)	0.3942*** (0.0668)	0.76 (0.0690)	0.00	0.23
SEq	0.1091*** (0.0361)	0.7196*** (0.0719)	0.2763*** (0.0719)	0.79 (0.0796)	0.00	0.38
UK	0.0000 (0.0306)	0.8745*** (0.0839)	0.1916 (0.0839)	0.70 (0.1427)	0.16	0.73

Notes: See footnote of Table 4 for a description.

Table A.11: Model (1), 2003–2007

Country	$\alpha_1$	$\alpha_2$	$\alpha_0$	$R^2$	Wald p	BG p
AT	0.0283 (0.0287)	0.5799*** (0.1111)	0.9019*** (0.2637)	0.34	0.00	0.12
BE	0.0277 (0.0248)	0.8361*** (0.0911)	0.3200* (0.1859)	0.73	0.11	0.55
DE	-0.0011 (0.0185)	0.8478*** (0.0437)	0.2186** (0.0885)	0.93	0.01	0.01
E.A	-0.0019 (0.0317)	0.9081*** (0.0460)	0.2264 (0.1379)	0.92	0.11	0.01
EL	-0.3007* (0.1696)	0.5523*** (0.1052)	4.1785*** (1.0516)	0.41	0.00	0.10
ES	0.0535 (0.0387)	0.7427*** (0.1123)	0.8993** (0.4309)	0.58	0.05	0.13
FI	0.0575*** (0.0194)	0.6918*** (0.0887)	0.5023*** (0.1489)	0.65	0.00	0.43
FR	0.0199 (0.0241)	0.6074*** (0.1370)	0.8857*** (0.3169)	0.41	0.01	0.42
IE	0.0803** (0.0354)	0.8175*** (0.0647)	0.3622* (0.1831)	0.83	0.08	0.02
IT	0.2631** (0.1002)	0.8462*** (0.0596)	-0.1120 (0.1620)	0.92	0.12	0.02
NL	0.2261* (0.1256)	0.8782*** (0.0507)	-0.0960 (0.2063)	0.90	0.42	0.01
SE	0.1558** (0.0657)	0.6863*** (0.0813)	0.2239** (0.0873)	0.80	0.01	0.45
SEq	0.1761** (0.0759)	0.6057*** (0.0982)	0.3561*** (0.1230)	0.70	0.00	0.49
UK	0.0492* (0.0281)	0.8249*** (0.0564)	0.1987** (0.0790)	0.82	0.01	0.32

Notes: See footnote of Table 4 for a description.

Table A.12: Model (2), 1993–2001

<i>Country</i>	$\alpha_1$	$\alpha_2$	$\alpha_0$	$R^2$	Wald p	BG p
AT	0.0363 (0.0270)	0.9387*** (0.0476)	0.0249 (0.0399)	0.92	0.49	0.20
BE	0.0693*** (0.0235)	0.7766*** (0.0592)	0.2396*** (0.0653)	0.86	0.00	0.02
DE	0.0473** (0.0218)	0.8496*** (0.0585)	0.1265** (0.0573)	0.90	0.03	0.91
E.A	0.0173 (0.0158)	0.9312*** (0.0415)	0.0880* (0.0475)	0.95	0.07	0.27
EL	0.0189** (0.0084)	0.8237*** (0.0487)	0.6467*** (0.1809)	0.84	0.00	0.65
ES	0.0486** (0.0186)	0.8387*** (0.0471)	0.2807*** (0.0952)	0.88	0.00	0.40
FI	0.0073 (0.0246)	0.9238*** (0.0405)	0.0996 (0.0631)	0.96	0.10	0.05
FR	0.0321* (0.0170)	0.8974*** (0.0452)	0.0898* (0.0476)	0.86	0.07	0.49
IE	0.0914*** (0.0237)	0.7854*** (0.0570)	0.3224*** (0.0937)	0.96	0.00	0.86
IT	0.0256 (0.0218)	0.9010*** (0.0495)	0.1659** (0.0732)	0.92	0.02	1.00
NL	0.0899*** (0.0219)	0.6688*** (0.0636)	0.4053*** (0.0805)	0.87	0.00	0.92
SE	0.1346*** (0.0394)	0.5760*** (0.0740)	0.3870*** (0.0722)	0.73	0.00	0.07
SEq	0.0829* (0.0423)	0.7569*** (0.0783)	0.2579*** (0.0838)	0.78	0.01	0.46
UK	-0.0277 (0.0311)	0.8721*** (0.0855)	0.2332 (0.1412)	0.70	0.08	0.73

Notes: See footnote of Table 5 for a description.

Table A.13: Model (2), 2003–2007

<i>Country</i>	$\alpha_1$	$\alpha_2$	$\alpha_0$	$R^2$	Wald p	BG p
AT	0.0305 (0.0278)	0.5725*** (0.1122)	0.9150*** (0.2555)	0.35	0.00	0.20
BE	0.0235 (0.0275)	0.8303*** (0.1001)	0.3415* (0.1959)	0.72	0.10	0.72
DE	-0.0171 (0.0150)	0.8253*** (0.0428)	0.2790*** (0.0816)	0.93	0.00	0.01
EA	0.0118 (0.0305)	0.9084*** (0.0457)	0.1966 (0.1375)	0.92	0.17	0.01
EL	-0.2552 (0.1692)	0.5559*** (0.1136)	4.0093*** (1.1475)	0.40	0.01	0.09
ES	0.0425 (0.0441)	0.7469*** (0.1171)	0.9151** (0.4412)	0.57	0.05	0.13
FI	0.0186 (0.0281)	0.7698*** (0.1095)	0.3994** (0.1799)	0.61	0.03	0.14
FR	0.0407* (0.0227)	0.5863*** (0.1361)	0.8939*** (0.3172)	0.43	0.01	0.17
IE	0.0651* (0.0346)	0.8225*** (0.0668)	0.3854** (0.1838)	0.82	0.05	0.01
IT	0.1738* (0.1001)	0.8796*** (0.0552)	-0.0233 (0.1780)	0.91	0.49	0.01
NL	0.2046 (0.1491)	0.8731*** (0.0455)	-0.0547 (0.2313)	0.90	0.60	0.00
SE	0.0450 (0.0406)	0.8004*** (0.0811)	0.2159** (0.1022)	0.77	0.02	0.11
SEq	0.0483 (0.0505)	0.7305*** (0.1061)	0.3420** (0.1451)	0.65	0.01	0.07
UK	0.0487 (0.0297)	0.8236*** (0.0614)	0.2017** (0.0802)	0.82	0.01	0.34

Notes: See footnote of Table 5 for a description.

Table A.14: Model (1), Cochrane-Orcutt estimates

<i>Country</i>	$\alpha_1$	$\alpha_2$	$\alpha_0$	$R^2$	Wald p
AT	0.0525*** (0.0184)	0.9224*** (0.0355)	0.0200 (0.0339)	0.98	0.40
BE	0.0354** (0.0153)	0.8691*** (0.0350)	0.1477*** (0.0418)	0.95	0.00
DE	0.0325* (0.0169)	0.9550*** (0.0344)	0.0172 (0.0442)	0.94	0.74
EA	0.0475*** (0.0139)	0.8742*** (0.0300)	0.1296*** (0.0385)	0.97	0.00
EL	0.0105 (0.0065)	0.8908*** (0.0355)	0.4015*** (0.1352)	0.95	0.00
ES	0.0618*** (0.0142)	0.8161*** (0.0321)	0.3008*** (0.0737)	0.96	0.00
FI	0.0346** (0.0135)	0.8725*** (0.0211)	0.1159*** (0.0235)	0.97	0.00
FR	0.0445*** (0.0147)	0.8478*** (0.0419)	0.1366*** (0.0460)	0.97	0.00
IE	0.0656*** (0.0134)	0.8554*** (0.0308)	0.2075*** (0.0547)	0.97	0.00
IT	0.0378** (0.0150)	0.9064*** (0.0298)	0.1146** (0.0522)	0.95	0.01
NL	0.0565*** (0.0168)	0.9008*** (0.0399)	0.0597 (0.0636)	0.95	0.23
SE	0.1353*** (0.0274)	0.6862*** (0.0511)	0.2340*** (0.0577)	0.79	0.00
SEq	0.1003*** (0.0285)	0.7994*** (0.0439)	0.1701*** (0.0556)	0.84	0.01
UK	0.0313 (0.0202)	0.8652*** (0.0441)	0.1653** (0.0635)	0.79	0.01

*Notes:* This table shows estimates of Model (1) using the iterated Cochrane-Orcutt method, 1993-2007. The country-specific sample periods are specified in Table A.1. The column *Wald p* reports the p-value of the Wald test of the restriction  $\alpha_1 + \alpha_2 = 1$ . Estimates of the indicator variable for the euro cash changeover are not reported. Heteroskedasticity robust standard errors in parentheses. \*, \*\* and \*\*\* indicate statistical significance at the 10%, 5% and 1% level, respectively.

Table A.15: Model (2), Cochrane-Orcutt estimates

<i>Country</i>	$\alpha_1$	$\alpha_2$	$\alpha_0$	$R^2$	Wald p
AT	0.0356* (0.0180)	0.9303*** (0.0366)	0.0365 (0.0335)	0.98	0.25
BE	0.0384*** (0.0130)	0.8599*** (0.0351)	0.1570*** (0.0441)	0.95	0.00
DE	0.0147 (0.0154)	0.9500*** (0.0343)	0.0454 (0.0454)	0.94	0.35
EA	0.0284** (0.0142)	0.9013*** (0.0329)	0.1191*** (0.0399)	0.97	0.00
EL	0.0082 (0.0066)	0.8933*** (0.0358)	0.4041*** (0.1356)	0.95	0.00
ES	0.0546*** (0.0160)	0.8201*** (0.0361)	0.3127*** (0.0767)	0.96	0.00
FI	0.0145 (0.0143)	0.8892*** (0.0218)	0.1335*** (0.0238)	0.97	0.00
FR	0.0450*** (0.0144)	0.8318*** (0.0412)	0.1564*** (0.0473)	0.96	0.00
IE	0.0640*** (0.0131)	0.8529*** (0.0311)	0.2201*** (0.0575)	0.96	0.00
IT	0.0223 (0.0153)	0.9256*** (0.0301)	0.1138** (0.0527)	0.95	0.02
NL	0.0548*** (0.0175)	0.9000*** (0.0397)	0.0666 (0.0643)	0.94	0.21
SE	0.0800*** (0.0277)	0.7541*** (0.0567)	0.2195*** (0.0607)	0.78	0.00
SEq	0.0603** (0.0298)	0.8437*** (0.0450)	0.1605*** (0.0548)	0.84	0.01
UK	0.0179 (0.0209)	0.8831*** (0.0458)	0.1541** (0.0639)	0.79	0.01

Notes: This table shows estimates of Model (2) using the iterated Cochrane-Orcutt method, 1993–2007. See footnote of Table A.14 for a description.

Table A.16: Model (1) in first differences

<i>Country</i>	$\alpha_1$	$\alpha_2$	$\alpha_0$	$R^2$	Wald p	BG p
AT	0.0837* (0.0463)	-0.1444 (0.0909)	0.0211 (0.0170)	0.06	0.00	0.02
BE	-0.0109 (0.0359)	-0.2553*** (0.0950)	-0.0049 (0.0115)	0.07	0.00	0.16
DE	0.0519* (0.0295)	0.1549 (0.1261)	0.0027 (0.0072)	0.06	0.00	0.04
EA	0.0971** (0.0374)	0.1000 (0.0761)	0.0001 (0.0057)	0.08	0.00	0.00
EL	0.1141*** (0.0432)	-0.2850*** (0.0914)	-0.0141 (0.0257)	0.11	0.00	0.02
ES	0.1194*** (0.0435)	-0.0664 (0.0815)	0.0017 (0.0164)	0.04	0.00	0.00
FI	0.1340*** (0.0441)	-0.1151 (0.1032)	0.0362** (0.0169)	0.13	0.00	0.67
FR	0.0311 (0.0296)	-0.0074 (0.0777)	0.0069 (0.0107)	0.01	0.00	0.00
IE	0.0719* (0.0381)	-0.2168*** (0.0798)	0.0143 (0.0174)	0.07	0.00	0.75
IT	0.2250*** (0.0493)	-0.1394* (0.0823)	-0.0025 (0.0146)	0.09	0.00	0.08
NL	0.0469 (0.0584)	-0.2905** (0.1412)	0.0066 (0.0123)	0.09	0.00	0.27
SE	0.1717*** (0.0511)	-0.1105 (0.1206)	-0.0080 (0.0270)	0.09	0.00	0.01
SEq	0.1512*** (0.0541)	-0.1594 (0.1080)	0.0016 (0.0242)	0.08	0.00	0.06
UK	0.0560 (0.0383)	-0.0465 (0.0908)	-0.0104 (0.0133)	0.03	0.00	0.56

*Notes:* This table shows OLS estimates of Model (1) in first differences, 1993-2007. The country-specific sample periods are specified in Table A.1. The column *Wald p* reports the p-value of the Wald test of the restriction  $\alpha_1 + \alpha_2 = 1$ . *BG p* is the p-value of the Breusch-Godfrey LM test statistic for first order residual correlation. Estimates of the indicator variable for the euro cash changeover are not reported. White standard errors allowing for heteroskedasticity in parentheses. \*, \*\* and \*\*\* indicate statistical significance at the 10%, 5% and 1% level, respectively.

Table A.17: Model (2) in first differences

<i>Country</i>	$\alpha_1$	$\alpha_2$	$\alpha_0$	$R^2$	Wald p	BG p
AT	0.0303 (0.0377)	-0.1818* (0.0965)	0.0222 (0.0179)	0.03	0.00	0.02
BE	0.0555*** (0.0191)	-0.2491*** (0.0834)	-0.0044 (0.0112)	0.11	0.00	0.29
DE	0.0196 (0.0277)	0.1276 (0.1153)	0.0035 (0.0077)	0.03	0.00	0.04
EA	0.0994*** (0.0326)	0.0116 (0.0775)	0.0001 (0.0060)	0.08	0.00	0.52
EL	0.0492 (0.0529)	-0.2896*** (0.0900)	-0.0202 (0.0265)	0.09	0.00	0.03
ES	0.1616*** (0.0354)	-0.1175 (0.0829)	0.0032 (0.0161)	0.07	0.00	0.44
FI	0.0670 (0.0414)	-0.1797 (0.1234)	0.0408** (0.0180)	0.05	0.00	0.02
FR	0.1090*** (0.0275)	-0.0321 (0.0760)	0.0078 (0.0103)	0.08	0.00	0.90
IE	0.0853* (0.0476)	-0.2236*** (0.0816)	0.0132 (0.0168)	0.07	0.00	0.64
IT	0.1329** (0.0598)	-0.1345 (0.0867)	-0.0041 (0.0149)	0.04	0.00	0.39
NL	0.0589 (0.0415)	-0.2980** (0.1426)	0.0064 (0.0121)	0.09	0.00	0.27
SE	0.1007 (0.0671)	-0.1441 (0.1184)	-0.0076 (0.0267)	0.04	0.00	0.06
SEq	0.0604 (0.0637)	-0.1913* (0.1102)	0.0044 (0.0239)	0.04	0.00	0.05
UK	-0.0099 (0.0445)	-0.0528 (0.0935)	-0.0112 (0.0139)	0.01	0.00	0.06

Notes: This table shows OLS estimates of Model (2) in first differences, 1993–2007. See footnote of Table A.16 for a description.



Table A.18: Near-rationality in Model (2), 1993–2007

<i>Country</i>	$\alpha_1$	$\alpha_2$	$\Delta\alpha_1$	$\Delta\alpha_2$	T	Median				
AT	-0.0067	(0.0278)	0.8823***	(0.0342)	0.0258	(0.0414)	-0.0212	(0.0446)	142	1.76
BE	0.0510	(0.0426)	0.8032***	(0.0575)	0.0012	(0.0438)	-0.0014	(0.0421)	175	1.88
DE	0.0187	(0.0187)	0.9414***	(0.0333)	0.0952**	(0.0405)	-0.0761**	(0.0351)	138	1.38
E.A	0.0657**	(0.0265)	0.8840***	(0.0381)	0.0514	(0.0339)	-0.0700*	(0.0400)	175	2.12
EL	0.0060	(0.0640)	0.8099***	(0.0530)	-0.0536	(0.0432)	0.0037	(0.0650)	175	3.81
ES	0.0525	(0.0371)	0.8009***	(0.0351)	0.0249	(0.0346)	-0.0443	(0.0424)	175	3.31
FI	0.0271	(0.0334)	0.8972***	(0.0304)	0.0145	(0.0391)	-0.0241	(0.0414)	141	1.36
FR	0.0325	(0.0221)	0.8289***	(0.0521)	0.0518	(0.0367)	-0.0537	(0.0338)	175	1.77
IE	0.0204	(0.0302)	0.8804***	(0.0362)	-0.0965**	(0.0425)	0.0872**	(0.0382)	139	2.65
IT	0.0339	(0.0507)	0.8815***	(0.0390)	0.0014	(0.0388)	-0.0170	(0.0539)	175	2.41
NL	0.0242	(0.0525)	0.8392***	(0.0441)	-0.0333	(0.0467)	0.0336	(0.0574)	175	1.82
SE	0.0807	(0.0550)	0.7178***	(0.0632)	0.0131	(0.0778)	-0.0477	(0.0786)	142	1.54
SE <sub>q</sub>	0.0782	(0.0546)	0.7645***	(0.0508)	0.0731	(0.0821)	-0.0937	(0.0893)	139	1.54
UK	-0.0765	(0.0512)	0.9290***	(0.0498)	-0.1412**	(0.0618)	0.1511**	(0.0586)	127	1.51

Notes: See footnote of Table 6 for a description.