Universität Hamburg

Department Economics and Politics

Perceived Inflation under Loss Aversion

Lena Dräger
Jan-Oliver Menz
Ulrich Fritsche

DEP Discussion Papers
Macroeconomics and Finance Series
5/2011

Hamburg, 2011

Perceived Inflation under Loss Aversion

Lena Dräger*
Jan-Oliver Menz**
Ulrich Fritsche‡

October 26, 2011

Abstract

Building on Prospect Theory, we apply the concept of loss aversion to the formation of inflation perceptions and test empirically for non-linearities in the inflation-perceptions relation for a panel of 10 Euro area countries. Specifically, under the assumption of loss aversion, inflation changes above a certain reference rate will be perceived more strongly. Rejecting rationality of inflation perceptions in general under symmetric loss and in a majority of cases under flexible loss functions, panel smooth transition models give evidence of non-linearities in the inflation perception formation regarding both actual inflation and time. This result is confirmed by dynamic fixed effects estimates, where the slope of the estimated value function is significantly steeper in the loss region and the implied average reference inflation rate is found close to 2%.

Keywords: Inflation Perceptions, Loss Aversion, Panel Smooth Transition Models, Dynamic Panel.

JEL classification: C33, D81, D82, E31.

---

*Corresponding author: University of Hamburg.
Email: Lena.Draeger@wiso.uni-hamburg.de
We would like to thank Jörg Breitung, Michael Lamla, Thomas Maag, Jan-Egbert Sturm as well as seminar participants at the University of Zurich, the University of Hamburg and the 2010 MMF Conference for their very helpful comments on earlier drafts of this paper and Ulrich Brandt, Torsten Osigus as well as Philipp Poppitz for excellent research support. All authors thankfully acknowledge financial support from the Deutsche Forschungsgemeinschaft.
**University of Hamburg, Email: Jan-Oliver.Menz@wiso.uni-hamburg.de
‡University of Hamburg & KOF Swiss Economic Instiute, ETH Zurich, Email: Ulrich.Fritsche@wiso.uni-hamburg.de

1 Introduction

When assessing macroeconomic models empirically, economists mostly use actual data as published by statistical institutes for the theoretical variables in these models. However, there exists overwhelming empirical evidence that peoples' knowledge and perception of these variables may deviate considerably from official statistical data and their underlying concepts, questioning the rationality of agents widely assumed since Muth (1961). Instead, inflation perceptions can be regarded as "nowcasts", as agents form beliefs over actual inflation on the basis of information currently available and potentially subject to biases.[1]

The gap between actual data and individuals' perceptions raises important policy questions. This is especially true for inflation. As argued by van der Klaauw et al. (2008), among others, if individuals have biased beliefs about inflation, this can seriously undermine the central bank's credibility. Conversely, a credible monetary regime can also influence inflation perceptions, for instance by creating a focal point at the inflation target.[2] Furthermore, relating to the concept of money illusion[3], the perception gap may lead to distortions in bargaining if individuals misperceive their actual real purchasing power. To assess the effectiveness of policy propositions suggested by macroeconomic models, it is thus necessary to understand how people form perceptions about macroeconomic variables and how these perceptions influence individual behavior.

So far, the literature on the formation of inflation perceptions has mainly focused on one stylized fact, namely the observed jump in perceptions after the Euro cash changeover in 2002, whereas actual inflation continued to stay on a low level. Explanations for this jump range from price intransparencies (Dziuda and Mastrobuoni, 2009), difficulties in applying the conversion rates (Ehrmann, 2006), a perceptual crisis (Eife, 2006, Eife and Coombs, 2007, Fullone et al., 2007 and Blinder and Krueger, 2004), macroeconomic illiteracy (Del Giovane et al., 2008, Cestari et al., 2008), a media bias (Lamla and Lein, 2008), and expectancy confirmation (Traut-Mattausch et al., 2004).

A number of papers furthermore analyze factors influencing perceived inflation in general. Del Giovane et al. (2008) design a detailed survey for Italian consumers in 2006. The authors report asymmetries in perceived inflation, since respondents stating that they have observed price decreases over the last five years report significantly lower inflation perceptions than those not

[1] See Blanchflower and Kelly (2008), Blinder and Krueger (2004), Jonung and Laidler (1988), Malgarini (2008), Curtin (2007) and van der Klaauw et al. (2008).

[2] Evidence for this channel has been found in inflation perception surveys for Sweden, see Bryan and Palmqvist (2005).

[3] See Fisher (1928) for the original contribution, and Shafir et al. (2004) and Fehr and Tyran (2007) for a Behavioral Economics perspective.

recalling any price decreases. Furthermore, survey responses suggest a strong impact of socioeconomic factors on inflation perceptions. This is in line with findings in Jonung (1981) who claims that inflation perceptions in Sweden differ significantly between genders. Furthermore, in a recent survey, Jonung and Conflitti (2008) report differences between age, gender, occupational and regional groups with respect to opinions of the Euro currency, which may also be reflected in inflation perceptions.

Lein and Maag (2011) analyze the formation of inflation perceptions for the EU and Sweden, using data from the Joint Harmonized EU Program of Business and Consumer Surveys and Sweden's Consumer Tendency Survey. The authors reject rationality of perceptions, since quantified inflation perceptions fail the rationality conditions of accuracy, unbiasedness and efficiency. They also find some evidence for the importance of frequently bought goods, and for the expectancy confirmation hypothesis in the Euro area after the cash changeover. This is in line with Döhring and Mordonu (2007), who report an influence of inflation expectations on perceptions in addition to actual inflation, estimating a dynamic panel model for the countries that adopted the Euro in 2002.

Following the deviation of perceived from actual inflation rates at the Euro cash changeover, Brachinger (2006, 2008) proposes an *Index of Perceived Inflation* (IPI) meant to capture movements in perceived inflation better than usual CPI inflation. The IPI index is constructed under the assumption that agents perceive inflation according to behavioral patterns defined in Prospect Theory by Kahneman and Tversky (1979) and Tversky and Kahneman (1981, 1991). These include the concepts of loss aversion with respect to above-average inflation and the availability bias. Jungermann et al. (2007) perform an experimental study of the assumptions underlying the IPI index, and find evidence of a loss aversion parameter of about 2.[4] However, their approach has been criticized by Hoffmann et al. (2006) for its use of arbitrary ad hoc assumptions.

This paper adds to the literature as follows. Building on Prospect Theory by Kahneman and Tversky (1979) and Tversky and Kahneman (1981, 1991), we empirically test for the existence of loss aversion with respect to a reference inflation rate affecting the formation of inflation perceptions.[5]

Developed as an alternative decision theory under risk and uncertainty opposed to traditional expected utility theory,[6] Prospect Theory proposes that individuals code price changes and evaluate them against a reference price,

[4] This relates well to studies of loss aversion in other areas, where approximately the same parameter has been found, see for example Tversky and Kahneman (1991), Hardie et al. (1993) and Rosenblatt-Wisch (2008).

[5] Note that this is one of the hypotheses underlying the construction of the IPI index in Brachinger (2006, 2008).

[6] See Starmer (2004) for an overview of developments in decision theory under risk.

where higher prices are perceived as losses and lower prices are perceived as gains. Since individuals display loss aversion, price increases are perceived more strongly than price decreases, the exact quantity being captured by the loss aversion parameter. Note that Prospect Theory defines loss aversion with respect to prices, while we analyze loss aversion with respect to inflation. This implies that a certain inflation rate is deemed 'normal', while inflation rates above the 'normal' rate are perceived as a loss. Also, analyzing loss aversion with respect to inflation entails a dynamic view of price developments, where reference prices and the reference inflation rate are grounded in households' historical experience (Malmendier and Nagel, 2011).[7]

Figure 1 shows a stylized value function describing the relation between perceived and actual inflation. The existence of loss aversion leads to a kink at the reference rate of inflation, with a steeper slope in the loss region where inflation rates are above the reference rate.[8]

< Figure 1 here >

We evaluate loss aversion regarding inflation by analyzing households' survey data on perceived inflation compiled by the European Commission in the *Joint Harmonized EU Program of Business and Consumer Surveys*. In order to apply for rationality tests, and to allow for the estimation of interpretable slope parameters, the qualitative survey answers are quantified with the method by Carlson and Parkin (1975), which has been adapted to a pentachotomous survey by Batchelor and Orr (1988). The sample then covers a panel of 10 Euro area countries from January 1996 to December 2010.

The empirical investigation takes the following route: First, we examine rationality criteria for inflation perceptions following the seminal work of Mincer and Zarnowitz (1969) and test for the unbiasedness and efficiency of perceived inflation. Underlying these rationality tests, however, is the assumption of a symmetric loss function, which is no longer appropriate if loss aversion with respect to high inflation rates is present. Thus, we extend the analysis by applying the quantile approach by Patton and Timmermann (2007), which accounts for both symmetric and asymmetric loss functions.[9]

Second, the existence and shape of a non-linear relation between perceived

[7]The concept of loss aversion has also been applied to other areas, such as brand choice or consumption patterns, see for instance Hardie et al. (1993), Camerer (2000), Rosenblatt-Wisch (2008), Foellmi et al. (2011) and Gaffeo et al. (2011).

[8]In order to determine the reference price, two routes can be followed. In the context of consumer choice, the reference price is given by the *fair* price, which is determined by consumers' perceptions of sellers' costs. This idea has first been proposed by Thaler (1985) and recently been pursued further by Rotemberg (2005, 2008). With regard to inflation perceptions, Brachinger (2006) argues that one could simply take a past price as the reference price. However, it is not clear whether one should use an average price of a bundle of goods and how long the reference time period should be.

[9]A similar approach is followed by Capistrán and Timmermann (2009) who analyze asym-

and actual inflation is investigated in a Panel Smooth Transition (PSTR) setting proposed by Gonzalez et al. (2005) and Fok et al. (2005). This approach allows the estimation of the reference inflation rate and the transition function, while accounting for potential structural breaks. Finally, we estimate dynamic fixed effects models with threshold variables, defining a time-varying reference inflation rate, as a robustness check and in order to determine the slope parameters and the location of the kink in a non-linear perception-inflation relation.

Our analysis suggests the following results: While we generally reject rationality of inflation perceptions in our sample, allowing for asymmetric loss functions yields a number of non-rejections in the second half of our sample period. This indicative result of asymmetries underlying the inflation-perceptions nexus is further confirmed by the PSTR models, which find non-linearities with respect to both actual inflation and time. Generally, results suggest a significantly stronger effect of actual inflation changes on perceptions once inflation is above a certain threshold, which is estimated to be in the range from 1.8% to 3.3%. Estimates from dynamic fixed effects models with a time-varying reference rate of inflation confirm this result and imply a reference rate close to 2%.

The remainder of the paper is structured as follows. Section 2 describes the data set, including the quantification method for the qualitative survey data and presents panel unit root and cointegration tests. Section 3 proceeds with presenting the estimation design, followed by a discussion of the results in Section 4. Finally, Section 5 concludes.

2 Data Set and Statistical Properties

2.1 Perceived and Actual Inflation

The hypothesis from Prospect Theory – there is a non-linear relationship between perceptions and inflation – is tested empirically for a panel of 10 EMU-Countries consisting of Austria, Belgium, Finland, France, Germany, Greece, Italy, the Netherlands, Portugal, and Spain for the time period from January 1996 to December 2010. Our sample thus covers the Euro area almost completely, and the sample period is long enough to enable us to test for possible structural breaks.

We use a quantified version of the balance statistic of Question 5 of the *Joint Harmonized EU Program of Business and Consumer Surveys* by the European Commission as our measure of perceived inflation. The officially published balance statistic of the survey provides only a qualitative measure from the pentachotomous survey, asking participants whether they think

metries regarding the formation of inflation expectations from the Survey of Professional Forecasters in the US.

prices have fallen/ stayed about the same/ increased at a slower rate/ increased at the same rate/ increased more rapidly over the last 12 months. Denoting the shares of answers in each category as s1, s2, s3, s4 and s5, the balance statistic is obtained as $s1 + 0.5s2 - 0.5s4 - s5$. While most empirical studies on perceived or expected inflation with data from the Joint Harmonized EU Program of Business and Consumer Surveys make use of the balance statistic, there exist methods to quantify the qualitative data. Given that we want to test the rationality of perceptions under symmetric or asymmetric loss functions explicitly, we have to rely on quantified perception data. We thus follow Döpke et al. (2008) and employ a version of the probability method proposed by Carlson and Parkin (1975) and modified by Batchelor and Orr (1988).[10]

The quantification method demands a scaling series that inflation perceptions are assumed to be based upon. We specifically assume that households observe the underlying medium-term trend of the true inflation rate correctly and proxy the current scaling value of inflation with a recursively estimated Hodrick-Prescott filter under the usual assumption of $\lambda = 14.400$ for monthly data. Business cycle fluctuations are therefore excluded from the expected value observed by households. Details are explained in the technical appendix of this paper (section A).[11]

Actual inflation rates are measured with annual inflation rates of harmonized consumer price indices (HICP) from Eurostat. All data are available on a monthly basis. Figure 2 shows the quantified perceptions together with the inflation rates for all countries of our sample. In most countries of our sample, on average quantified perceptions track actual inflation relatively closely. Nevertheless, significant deviations do occur, for instance at the Euro cash changeover or at the spike in inflation before the financial crisis.

< Figure 2 here >

2.2 Unit Roots and Cointegration

We test the time series of actual and perceived inflation rates for panel unit roots in order to avoid spurious regressions. Details on the test statistics and results are given in the appendix (section B). In line with results in the literature, e.g. Lein and Maag (2011), we find that the null hypothesis of a unit root in inflation is mostly rejected, while perceptions seem more persistent. Generally, empirical evidence on the order of integration of inflation series is mixed, Altissimo et al. (2006) conclude in a survey that empirical findings

[10] See also Nielsen (2003) for a survey.

[11] See Nardo (2003) for a critical overview on quantification of survey data. Maag (2009) analyses both quantitative and qualitative measures of Swedish inflation perceptions and expectations. The author finds that quantified data and balance statistics are of equal accuracy and that both are highly correlated with the mean of actual quantitative beliefs.

seem to lean towards stationarity of inflation. Due to the mixed results with respect to stationarity of perceived and actual inflation rates, we proceed to test for panel cointegration between perceptions and inflation.
A detailed description of the panel cointegration tests applied and of the results is again given in the appendix. We find significant evidence for cointegration: For the whole sample period, all tests reject the null of no cointegration at the 1% level. This, quite intuitive, result is in line with findings in Lein and Maag (2011) for a similar sample. Considering the results from both panel unit root and cointegration tests, we estimate regressions in the analysis in levels, making use of the super-consistency argument by Engle and Granger (1987).

3 Estimation Design

3.1 Non-Rationality of Perceived Inflation under Flexible Loss Functions

The paper of Mincer and Zarnowitz (1969) on the empirical investigation of the rationality of professional forecasts initiated a bulk of literature on the econometrics of rationality tests. These tests are regularly applied to inflation forecasts of professional forecasters and households. Under the assumption that aggregate information dissipates slowly throughout the economy (i.e. "sticky information"), households' perceptions can be seen as forecasts made today (so-called "nowcasts"), based on all the available information in the current time period. Therefore, most rationality tests in the forecast evaluation literature can be applied to inflation perceptions as well. However, rationality tests typically rely – implicitly or explicitly – on the axiom of a symmetric loss function (see Granger, 1969; Bachelor and Peel, 1998; Patton and Timmermann, 2007). The first class of such tests dates back to the seminal paper of Mincer and Zarnowitz (1969) and is therefore called a "Mincer-Zarnowitz regression". The test is constructed as a joint test of the unbiasedness and efficiency of the forecast. Dufour (1981) and Campbell and Ghysels (1995) suggest non-parametric tests on both aspects based on a usual Wilcoxon-Sign-Rank-test (see Wilcoxon, 1945).[12].
The symmetric loss function might be a good approximation for a wide range of cases. However, asymmetric loss functions are also quite plausible and especially suitable in cases where a loss aversion model is to be tested. To test for rationality in a broader sense, we therefore use a more flexible approach and employ an indicator function (or quantile approach) as proposed by Patton and Timmermann (2007). Under the joint assumption that the relevant loss is a function solely of the forecast error and that the functional form of

[12]We report results of non-parametric rationality tests based on a symmetric loss function in the appendix (section C)

the loss function is homogenous in the forecast error, we can use an indicator function which takes the value of one if the forecast is equal to or larger than the realization (i.e. the perception is higher than the true inflation rate) and estimate the following model:

$$I_t = \alpha_0 + \alpha_1 \hat{y_t}|t \tag{1}$$

where $\hat{y_t}|t$ is the perception in time t. Patton and Timmermann (2007) show that the indicator variable I_t should be independent of any element in the information set and therefore the restriction $\alpha_0 = \alpha_1 = 0$ should hold. We report results from a probit model which is satisfied due to the binary nature of the data.[13]

3.2 A Panel Smooth Transition Approach

Having tested for asymmetries in the inflation-perceptions relation in the previous section, we proceed to investigate the (potentially nonlinear) relationship further by applying the Panel Smooth Transition (PSTR) model developed by Gonzalez et al. (2005) and Fok et al. (2005). Recently, this approach has been used in a number of applications. With its help, Hurlin and coauthors investigate the nonlinear relationship between public capital and output (Colletaz and Hurlin, 2006), the Feldstein-Horioka puzzle (Fouquau et al., 2008), and energy demand (Destais et al., 2009). Others have used the PSTR-model to analyze nonlinearities with respect to health care expenditure and GDP (Chakroun, 2010, Mehrara et al., 2010), the effects of financial development and growth (Jude, 2010), and the link between inflation and growth (Ibarra and Trupkin, 2011).

The increasing popularity of the PSTR-model might be due to the fact that it has two advantages over simple fixed effects estimations.

First, the model allows to explicitly test for nonlinearity, and to endogenously determine both the threshold and the degree of nonlinearity. Second, as will become clear below, the PSTR-model also allows for different coefficients of the explanatory variables over time and over cross-section units, whereas the dynamic fixed effects model only captures panel heterogeneity by fixed individual and time effects. One caveat applies, however. By now, it is still unclear how the PSTR-model, and also the PTR-model proposed by Hansen (1999), behaves with respect to dynamic panels including lags of the endogenous variable. Given the high serial correlation of inflation perceptions, this problem deserves particular attention in our setting. Hence, we will present PSTR estimates both including and excluding lagged inflation perceptions from the nonlinear part, and also apply dynamic fixed effects. Moreover,

[13] We experimented a bit with the functional form of the underlying cumulative distribution. In most cases, logit models yield qualitatively similar results. Results are available from the authors upon request.

we estimate single country regressions using the Smooth Transition Autoregressive (STAR)-model developed by Granger and Teräsvirta (1993) which is able to deal with the existence of lagged endogenous variables.

Specifying and Estimating the Model

The PSTR-model, allowing for nonlinearities between the explanatory variables and the dependent variable, and a smooth transition between different regimes separated via a transition variable, is defined as:

$$y_{it} = \mu_i + \beta_0 x'_{it} + \beta_1 x'_{it} G(q_{it}; \gamma, c) + u_{it} \tag{2}$$

where $t = 1, \ldots, T$ and $i = 1, \ldots, N$ denote the time and the cross-section dimension, respectively, and μ_i captures the fixed individual effects. For G, one uses the logistic function

$$G(q_{it}; \gamma, c) = \left(1 + exp\left(-\gamma_G \prod_{j=1}^{m} (q_{it} - c_j)\right)\right)^{-1}, \tag{3}$$

Here, q_{it} is the transition variable: in case we find a nonlinear relationship, the coefficients of the explanatory variables change in line with the value of the transition variable. $c = (c_1, \ldots, c_m)'$ is an m-dimensional vector of *location parameters*, i.e. the *number* of thresholds and regimes. For example, if $m = 1$, we have one threshold, and two regimes, whereas for $m = 2$, the model consists of three regimes, a middle one, and two identical outer regimes. Finally, γ defines the steepness of the transition function, i.e., with $\gamma = 0$ we are back to the linear model, and with $\gamma \rightarrow \infty$, the model tends to a regime-switching model as developed by Hansen (1999).

This model can be understood in two ways. First, it can be interpreted as a regime-switching model, in which we get one coefficient β_0 in regime one, and another coefficient $\beta_0 + \beta_1$ if we are in regime two. Observations are divided into regimes via the transition variable q_{it}, and the transition between regimes might be smooth or immediate, depending on the value of γ. However, the model can also be seen as allowing for a large number of small regimes, each one being determined by a specific value of the transition variable. Applied to our research question: If individuals consider inflation differently depending on whether it is above or below a certain threshold, we will have a two-regime model, high- and low-inflation. By contrast, we could also interpret the findings in such a way that individuals adjust smoothly to changes in the inflation rate, resulting in various inflation/perception-regimes.

The PSTR-model can also be extended to allow for more than one transition function, in order to capture larger degrees of nonlinearity:

$$y_{it} = \mu_i + \beta_0 x'_{it} + \sum_{j=1}^{r} \beta_j x'_{it} G_j \left(q_{it}^{(j)}; \gamma_j, c_j \right) + u_{it} \tag{4}$$

where r is the number of transition functions. If $r = 0$, we are back to the linear model. Moreover, this latter specification can be used to test for a (gradual or immediate) structural break.[14] This is done by using a second transition function

$$T(t^*; \gamma, c) = \left(1 + exp \left(-\gamma_T \prod_{j=1}^{h} (t^* - c_j) \right) \right)^{-1} \tag{5}$$

with the time index $t^* = t/T$ as additional transition variable.

Summing up, this model allows for a number of possible relationships between the variables of interest. Applied to our research question regarding the link between inflation perceptions and the actual inflation rate, we would expect a model with $m = 1$ and $r = 1$, i.e. a model with one threshold (the reference inflation rate), two regimes (one above and one below the reference rate), and with one transition function. In case we find a structural break, we might get a model with r=2 and two different transition variables, the inflation rate and the time dimension.

Before estimating the model, we have to determine the number of location parameters m, and the number of transition functions.[15]

We start by testing $H_0 : r = 0$ vs. $H_1 : r = 1$: If we do not reject the null hypothesis, we conclude that the relationship is in fact linear, and estimate a dynamic fixed effects model. If we reject the null, we continue with testing $H_0 : r = 1$ vs. $H_1 : r = 2$, i.e., we test for remaining non-linearity. We repeat this procedure until we cannot reject the null hypothesis anymore, ending up with the optimal number of transition functions. The tests are carried out by replacing $G(q_{it}; \gamma, c)$ in (2) by its first-order Taylor expansion around $\gamma = 0$. Hence, one estimates the auxiliary regression

$$y_{it} = \mu_i + \beta_0^* x'_{it} + \beta_1^* x'_{it} q_{it} + \beta_2^* x'_{it} q_{it}^2 + \ldots + \beta_m^* x'_{it} q_{it}^m + u_{it}^* \tag{6}$$

where the vectors $\beta_1^*, \ldots, \beta_m^*$ are multiples of γ. Thus, testing $H_0 : \beta_1^* = \ldots = \beta_m^* = 0$ is equivalent to testing $H_0 : \gamma = 0$ in (2), in which case the model collapses into a standard linear fixed effects panel regression.

Next, estimation is carried out in three steps.[16] First, the individual effects μ_i are removed by subtracting the individual-specific means. Then, initial values for γ_j and c_j are chosen by means of a grid search, and, given

[14] See Gonzalez et al. (2005).

[15] See for details Colletaz and Hurlin (2006).

[16] We use the RATS code developed by Gilbert Colletaz and Christophe Hurlin that the authors have kindly made available online, see http://www.univ-orleans.fr/deg/masters/ESA/GC/gcolletaz_R.htm.

these values, the coefficients β_j are estimated with OLS. Third, using these estimates, γ_j and c_j are estimated by nonlinear least squares, allowing to calculate the final estimates for β_j. After the estimations, information criteria are compared in order to determine the optimal number of location parameters.

Concerning the interpretation of the estimated parameters, we can calculate the partial derivatives as

$$e_{it} = \frac{\partial y_{it}}{\partial x_{it}} = \beta_0 + \sum_{j=1}^{r} \beta_j G(q_{it}; \gamma_j, c_j) \tag{7}$$

or, if the threshold variable is the same as the explanatory variable, as:

$$e_{it} = \frac{\partial y_{it}}{\partial x_{it}} = \beta_0 + \sum_{j=1}^{r} \beta_j G_j(q_{it}; \gamma_j, c_j) + \sum_{j=1}^{r} \beta_j x_{it} \frac{\partial G_j(q_{it}; \gamma_j, c_j)}{\partial x_{it}} \tag{8}$$

It is important to note that these derivatives cannot directly be interpreted as elasticities, given that their value depends on $G(q_{it}; \gamma, c)$. More precisely, if the transition function tends either to 0 or to 1, we can determine the elasticities in the extreme regimes, β_0 and $\beta_0 + \beta_1$. However, given that the transition between regimes might be smooth, we receive a number of elasticities defined as a weighted average of the extreme values. This means that the PSTR-model allows for different coefficients of the explanatory variables for each country over time: In each period, and for each individual country, the transition function takes on a different value resulting in a specific value for the elasticity.

Besides these time-varying and individual-specific coefficients, we can interpret the signs of the estimated parameters. If $\beta_1 G$ is found to be positive, this means that the estimated effect from inflation on perceptions rises with the inflation rate, i.e., the higher the level of the inflation rate, the stronger the impact of a one percentage point increase on inflation perceptions.

Estimated Models

Using the inflation rate and time as transition variables, and the entire data set 1996m01–2010m12, we first test for nonlinearity with respect to both transition variables. This takes into account the structural break between inflation perceptions and inflation around the Euro cash changeover in January 2002, which has been documented by a number of researchers.

Hence, in case we find nonlinearity, we estimate the following model:

$$\begin{aligned} \pi^p_{i,t} &= \mu_i + \alpha_0 \pi^p_{i,t-1} + \beta_0 \pi_{i,t} + \beta_1 \pi_{i,t} G(\gamma_1, c_{m,1}, \pi_{i,t}) \\ &\quad + \beta_2 \pi_{i,t} T(\gamma_2, c_{m,2}, t^*) + \varepsilon_{i,t} \end{aligned} \tag{9}$$

For the sake of comparison, we re-estimate the model using lagged inflation perceptions also in the non-linear part.

$$\begin{aligned}\pi^p_{i,t} &= \mu_i + \alpha_0\pi^p_{i,t-1} + \beta_0\pi_{i,t} + \left[\alpha_1\pi^p_{i,t-1} + \beta_1\pi_{i,t}\right] G(\gamma_1, c_{m,1}, \pi_{i,t}) \\ &+ \left[\alpha_2\pi^p_{i,t-1} + \beta_2\pi_{i,t}\right] T\left(\gamma_2, c_{m,2}, t^*\right) + \varepsilon_{i,t} \qquad (10)\end{aligned}$$

Depending on whether we find one or two regimes ($m = 1$, or $m = 2$), the transition functions become:

$$G(c_1)/T(c_1) = (1 + exp(-\gamma_1(q_{it} - c_1)))^{-1} \qquad (11)$$

$$G(c_1, c_2)/T(c_1, c_2) = (1 + exp(-\gamma_1(q_{it} - c_1) \cdot (q_{it} - c_2)))^{-1} \qquad (12)$$

3.3 Dynamic Fixed Effects

After testing for a non-linear relation between perceived and actual inflation in a general panel smooth transition setting, we estimate the change in the slope of the value function, as well as the change in the intercept, in a dynamic panel fixed effects model. Note that applying the well-known dynamic panel estimator to our model also allows us to check for robustness of the PSTR results, where so far applicability to models including a lagged endogenous term has not been thoroughly investigated.
Assuming that the transition from the "gain" to the "loss" region takes the form of a kink as in Figure 1, we construct two threshold-dummies that serve to capture the periods where losses in the form of rising inflation occurred. If the hypothesis of loss aversion holds, we should find a significantly stronger impact of those "loss" periods on perceived inflation than of the "gain" periods in inflation. The threshold-dummies for all $i = 1, 2, ..., 10$ countries in the panel are defined as follows:

$$\text{thold}_{1,it} = \begin{cases} 1 \text{ if } \pi_{it} > \pi^{MA}_{it} \\ 0 \text{ otherwise,} \end{cases}$$

$$\text{thold}_{2,it} = \begin{cases} 1 \text{ if } \pi_{it} > \pi^{HP}_{it} \\ 0 \text{ otherwise,} \end{cases}$$

where π^{MA}_{it} represents a 13-months backward-looking moving-average of inflation and π^{HP}_{it} stands for recursively HP-filtered inflation. We thus assume that the medium-term trend in inflation is observed correctly and serves as the time-varying reference inflation rate. This is in line both with our assumption for the quantification procedure and with the theoretical argument that loss aversion regarding inflation implies a dynamic view on the perception of prices and, thus, calls for a reference rate grounded in historical experience. The threshold dummies take on the value of one for periods with above-average inflation, and zero otherwise.

The threshold-dummies are then included in a dynamic fixed effects model of inflation perceptions, both individually and combined with HICP inflation rates. Thereby, we can account both for a change in the slope parameter and for a change in the intercept during periods with above-average inflation rates. We thus estimate the following model:[17]

$$\pi^p_{it} = \alpha_0 + \alpha_1 thold_{1,2it} + \alpha_2 \pi^p_{it-1} + \beta_1 \pi_{it} + \beta_2(\pi_{it} * thold_{1,2it}) + \gamma_{i0} + \varepsilon_{it} \quad (13)$$

A significantly positive coefficient β_2 in equation (13) suggests higher perceived inflation rates in periods of above-average inflation for our panel and, thus, gives evidence of loss aversion with respect to inflation. Regarding the change in the intercept captured by α_1, if we expect the reference rate of inflation to be positive, the kink should lie above the origin and the intercept for "loss" periods α_1 would be below the normal intercept due to the steeper slope of the value function in the loss region. In the presence of loss aversion, we thus expect β_2 to be significantly positive and α_1 significantly negative. Note that equation (13) models loss aversion with respect to inflation as a long-run phenomenon, in line with the theory in Kahneman and Tversky (1979).

4 Results

4.1 Rationality Tests under Flexible Loss Functions

Evaluating rationality of inflation perceptions under a wide range of loss functions, we report results of the quantile test for rationality introduced by Patton and Timmermann (2007).[18] The results in Table 1 indicate that even under the mild assumptions of the test used here, the null of rationality has to be rejected in almost all cases if we test over the full sample period.
However, visual inspection as well as the results of formal structural break tests reported later in the paper, lead us to conclude that at least one or even two sample splits might be necessary to control for breaks. Interestingly, the results change to some extent once we control for structural breaks either around the Euro cash changeover or around the spike in inflation rates shortly before the financial crisis turmoil. Especially if we control for the cash changeover break, the number of rejections drop significantly. This is in contrast to the well-known results from traditional Mincer-Zarnowitz regressions and other rationality tests based on the symmetric loss assumption, which

[17] We tested for possible endogeneity of inflation rates in equation (13), but found no correlation between π_{it} and ε_{it} in any of the specifications.

[18] Results of non-parametric rationality tests under the symmetric loss assumption as in Campbell and Ghysels (1995) are given in the appendix of this paper in section C. Further results of tests in the spirit of Mincer and Zarnowitz (1969) are available from the authors upon request.

uniformly reject the rationality of inflation perceptions (Jonung and Laidler, 1988; Lein and Maag, 2011; Dräger, 2011). We interpret this as a hint that the result of "non-rationality" observed in other studies might to some extent be driven by asymmtric loss functions, due for instance to loss aversion.

< Table 1 here >

4.2 Panel Smooth Transition Regression

Next, we turn to the results of the Panel Smooth Transition Model. Beginning with the linearity tests shown in Table 2, we note that linearity with respect to inflation is rejected if we allow for two thresholds (m=2), but not for models with only one threshold. However, if we turn to test the null hypothesis of one transition function (r=1) against the alternative of at least two nonlinear functions, the null is also rejected for a model with one threshold.[19] This seemingly contradictory result might be due to the structural break around the Euro cash changeover (see below), a fact that is supported by the strong rejection of linearity with respect to time. Hence, we continue to estimate models with one transition function for each transition variable $inflation$ and $time$.

< Table 2 here >

Next, we have to decide on the optimal number of thresholds, m. Table 3 displays the information criteria and the residual sum of squares (RSS) for different specifications. We use time as second transition variable, and distinguish between models with one and two thresholds (1,1 and 2,2), in addition to only one threshold for inflation and two thresholds for time (1,2). The results are fairly clear-cut. Regarding the model with only inflation in the nonlinear part, both the AIC and the Schwarz criterion prefer the model with $m = (1, 2)$. By contrast, allowing for lagged inflation in the nonlinear part, the information criteria choose $m = (1, 1)$ as the best specification. Hence, based on the test results, we continue to estimate $(1, 1)$- and $(1, 2)$-models for both linear and nonlinear lagged perceptions.[20]

< Table 3 here >

The resulting parameter estimates are given in Table 4. As it turns out, the results are fairly similar across the different specifications. With regard

[19] Restricting lagged perceptions to be linear does not allow to test for more than one transition function.

[20] As noted by Gonzalez et al. (2005), the test results have to be interpreted with caution, since the results are influenced by cross-country heteroscedasticity which cannot be accounted for yet. Hence, we do not a priori refrain from estimating the $(1, 1)$-model, however, we are confident to exclude the $(2, 2)$-specification given that in this case, the convergence of the algorithm is largely dependent on the starting values of the grid search.

to time, we always find large $\gamma_T{'}s$ pointing to the existence of at least one structural break. The associated thresholds 0.4 and 0.8 roughly belong to January 2002, the date of the Euro cash changeover, and to July 2008, the beginning of the financial crisis. Concerning the steepness of the transition function using inflation as threshold variable, we get $\gamma_G{'}s$ between 0.8 and 2.2, however, the two lowest values estimated for the model using nonlinear lagged perceptions are not significantly different from zero. The associated threshold inflation rates range from 1.8% to 3.3%. Next, we find inflation perceptions to be quite persistent with estimates of their lagged value close to 0.9.

Regarding the effect from actual inflation on perceptions, we thus do find support for loss aversion with respect to high inflation rates: All nonlinear coefficients $\pi_t G$ are significantly positive, meaning that individuals perceive changes in inflation more strongly if inflation is above a certain threshold. Moreover, the coefficient with respect to time $\pi_t T$ suggests that this effect either increased after the Euro cash changeover, or that the effects have been lower before the Euro cash changeover and after the beginning of the financial crisis.

< Table 4 here >

Note that these findings are also supported by the individual STAR regressions that we have estimated as robustness checks for the time span 2002m1–2010m12.[21] The estimated thresholds lie between 2.5% and 3.5%, with the exception of Spain where we estimate a threshold of 4%. Moreover, in case of Austria, Belgium, Finland, France, Germany, Italy, and Portugal, we find rather large $\gamma_G{'}s$, and positive coefficients for inflation in the nonlinear part.

Finally, we illustrate the differences in the effects of actual inflation on perceived inflation over time and between countries. For that purpose, Figure 3 shows scatter plots of actual inflation and the country- and time-specific elasticities computed according to equation (8) for the PSTR-model allowing only for the structural break at the Euro cash changeover.[22] The black line shows the relationship after the Euro cash changeover, and the gray line the link for the period before. For all countries, we can observe a nonlinear, smooth relationship between actual inflation and the elasticities. Besides, the figures show an upward shift in the turning point of the transition function in Austria, France, and Germany; a hint that in those countries, the reference rate of inflation might have increased after the Euro cash changeover.

< Figure 3 here >

[21] The detailed results are available upon request. Prior to the Euro introduction, the results are not reliable due to the short sample period.

[22] Allowing for the second structural break at the time of the financial crisis yields qualitatively similar results.

4.3 Dynamic Fixed Effects Estimations of Loss Aversion

Finally, we present results of the panel estimation of loss aversion as in equation (13) with both threshold dummies in Table 5, where the first column additionally reports a test for adaptive perceptions.

$<$ Table 5 here $>$

Due to our finding of cointegration between actual and perceived inflation, we estimate all equations in levels, using dynamic fixed effects to account for the high degree of persistence in perceived inflation. Even if this estimator suffers from the Nickell (1981) bias, this is not a severe problem since T is significantly larger than N in our sample. For the same reason, the Arellano and Bond (1991) estimator would be computationally inefficient. Hence, we employ the dynamic fixed effects estimator and check for a possible influence from the Nickell bias by using the 'Least Squares Dummy Variable Corrected' (LSDVC) estimator proposed by Bruno (2005).[23] The results are robust across all estimators. Additionally, conducting the Pesaran (2004) and the Breusch and Pagan (1980) tests of error cross-section dependence reveals that residuals are correlated across panels.[24] Hence, we present estimates with correlated panels corrected standard errors.

Estimation results from the PSTR model suggest the existence of two structural breaks in the perception-inflation relation over our sample period, namely at the Euro cash changeover in January 2002 and at the onset of the financial crisis in July 2008. Therefore, we conduct Quandt-Likelihood-Ratio tests for single country estimations of the model in (13) with both thresholds. The test runs individual structural break tests over each month in the full sample period and selects the date with the maximum Wald F-Statistic as the break date. Results presented in Table D.1 in the appendix imply that, over the full sample period, the structural break during the months preceding the financial crisis dominates over the break at the Euro introduction.[25]

We thus estimate the model in equation (13) over the full sample period 1996m1–2010m12 and compare the results to estimates from a restricted sample period 1996m1–2007m12, excluding the dominant break.[26]

Throughout all models, we find that lagged inflation perceptions yield a highly significant coefficient of about 0.94–0.96. This suggests indeed a high

[23] Results are available from the authors upon request.

[24] Test results are available from the authors upon request.

[25] Restricting the sample period to 1996m1–2007m12, Quandt-Likelihood-Ratio tests identify the second break at the Euro cash changeover in almost all sample countries. Results are available from the author upon request.

[26] We further estimated models accounting for the break in January 2002 by including a dummy variable taking on the value of 1 for the months 2001m11–2002m2. However, the dummy was insignificant in all models and the coefficients remained robust. Estimation results are available upon request.

degree of persistence in inflation perceptions; a results which was also implied by the panel unit root tests above. We thus cannot rule out that inflation perceptions in our panel are to some degree formed adaptively. In the preliminary model without threshold dummies, the coefficient on actual inflation implies a long-run impact of 1.2 from actual to perceived inflation. Accounting for a possible effect of loss aversion on the perception-inflation relation, both models for the full sample period yield a significant coefficient β_2, indicating that inflation indeed influenced perceptions significantly stronger in periods with above-average inflation rates. As expected, both models also find a negative coefficient α_1 for the intercept dummy $thold_{1,2}$. However, only the model with $thold_2$ constructed with recursively HP-filtered inflation yields a significantly negative α_1 and β_1 significant at the 1% level. Both the Akaike and the Bayes information criterion also prefer the second model.
Comparing results over the full sample period to those from the restricted sample 1996m1–2007m12, we find that the coefficients α_1 and β_2 remain largely unchanged, while the coefficient β_1 measuring the overall impact of inflation on perceptions is reduced from about 0.06 to about 0.05–0.04. Simultaneously, inflation perceptions seem to be more persistent when disregarding the turbulent recent years. Overall, results suggest that loss aversion with respect to inflation is a persistent long-run phenomenon, while the general impact of actual on perceived inflation may be reduced in periods of stable inflation rates.

< Figure 4 here >

In order to visualize the results given in Table 5, we simulate the value function regarding inflation implied by the long-run coefficients of the dynamic fixed effects model with $thold_2$ over the full estimation period. The resulting value function is depicted in Figure 4 for actual inflation rates ranging from 0% to 4%. The non-linear value function under loss aversion is marked by the thick line, whereas the dashed lines indicate the non-applicable linear parts in the gain and loss regions, respectively. In line with the stylized function depicted in Figure 1, the slope of our estimated value function is significantly steeper in the loss region with inflation rates above the reference rate. The average reference rate of inflation at the kink implied by the long-run coefficients of the model is found close to 2%, which is in line with our results from the PSTR-models and which coincides with the implicit inflation target by the ECB.

5 Conclusion

This paper investigates whether the concept of loss aversion from Prospect Theory by Kahneman and Tversky (1979) and Tversky and Kahneman (1981,

1991) can be meaningfully applied to provide explanations for individuals' formation of inflation perceptions. Analyzing a panel of 10 Euro area countries for the sample period from January 1996 to December 2010, we find some evidence for asymmetries regarding the formation of inflation perceptions.

First, in line with other findings in the literature (e.g. Jonung and Laidler, 1988; Lein and Maag, 2011), we reject rationality of inflation perceptions over the full sample period under the flexible quantile approach by Patton and Timmermann (2007), which allows for both symmetric and asymmetric loss functions. However, once the possibility of a structural break at the Euro cash changeover is accounted for, rationality can no longer be rejected in a number of countries, while traditional rationality tests under symmetric loss functions continue to reject rationality. This result might be driven by the effect of asymmetric loss functions, for instance due to loss aversion.

Second, we investigate the degree and type of non-linearity in the perception-inflation nexus by estimating a panel smooth transition (PSTR) model proposed in Gonzalez et al. (2005) and Fok et al. (2005). Estimating models with two transition functions regarding actual inflation and time, the results suggest reference inflation rates in the range from 1.8% to 3.3% and either one or two structural breaks, one at the Euro cash changeover and one at the beginning of the financial crisis. Regardless of the specification, all PSTR models find a significantly stronger effect of changes in actual inflation on perceptions in the upper inflation regime, with an increase in coefficients between 0.4–0.6 once inflation is above the estimated reference rate.

Finally, since the applicability of PSTR models to dynamic panels including a lagged endogenous variable has not been fully evaluated, results from a dynamic fixed effects model with time-varying reference rates of inflation are presented as a check for robustness. Both for the full sample period, and for a shorter time span excluding the dominant structural break before the financial crisis, our results suggest that inflation rates above the threshold are indeed perceived significantly stronger. A simulated value function describing the perception-inflation nexus with the long-run coefficients implied by the model has a steeper slope in the loss region and an average reference inflation rate at about 2%.

Overall, we thus find some support of a non-linear relation between actual and perceived inflation rates in our EMU sample. While this leads to the rejection of standard rationality tests, the concept of loss aversion from Prospect Theory seems to describe the non-linear perception-inflation nexus relatively well with an implied average reference rate close to the implicit inflation target by the ECB.

References

Altissimo, F., M. Ehrmann, and F. Smets (2006). Inflation Persistence and Price-Setting Behaviour in the Euro Area – A Summary of the IPN Evidence. *ECB Occasional Paper 46*.

Arellano, M. and S. Bond (1991). Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations. *Review of Economic Studies 58*(2), 277–297.

Bachelor, R. and D. A. Peel (1998). Rationality Testing under Asymmetric Loss. *Economics Letters 61*, 49–54.

Batchelor, R. A. and A. B. Orr (1988). Inflation Expectations Revisited. *Economica 55*(219), 317–331.

Blanchflower, D. G. and R. Kelly (2008). Macroeconomic Literacy, Numeracy and the Implications for Monetary Policy. *Working Paper Bank of England*.

Blinder, A. and A. Krueger (2004). What Does the Public Know about Economic Policy, and How Does It Know It? *Brookings Paper on Economic Activity* (1), 327–387.

Brachinger, H. W. (2006). Euro or "Teuro"?: The Euro-Induced Perceived Inflation in Germany. *University of Fribourg Switzerland Department of Quantitative Economics Working Paper 5*.

Brachinger, H. W. (2008). A New Index of Perceived Inflation: Assumptions, Method, and Application to Germany. *Journal of Economic Psychology 29*(4), 433–457.

Breusch, T. and A. R. Pagan (1980). The Lagrange Multiplier Test and its Applications to Model Specification in Econometrics. *Review of Economic Studies 47*(1), 239–253.

Bruno, G. S. (2005). Approximating the Bias of the LSDV Estimator for Dynamic Unbalanced Panel Data Models. *Economics Letters 87*(3), 361–366.

Bryan, M. and S. Palmqvist (2005). Testing Near-Rationality Using Detailed Survey Data. *Federal Reserve Bank of Cleveland Working Paper Series 05-02*.

Camerer, C. F. (2000). Prospect Theory in the Wild: Evidence from the Field. In D. Kahneman and A. Tversky (Eds.), *Choices, Values and Frames*, Chapter 16, pp. 288–300. Cambridge, UK: Cambridge University Press.

Campbell, B. and E. Ghysels (1995). Federal Budget Projections: A Nonparametric Assessment of Bias and Efficiency. *The Review of Economics and Statistics 77*(1), 17 – 31.

Capistrán, C. and A. Timmermann (2009). Disagreement and Biases in Inflation Expectations. *Journal of Money, Credit and Banking 41*(2–3), 365–396.

Carlson, J. A. and M. Parkin (1975). Inflation Expectations. *Economica 42*(166), 123–138.

Cestari, V., P. Del Giovane, and C. Rossi-Arnaud (2008). Memory for Prices and the Euro Cash Changeover: An Analysis for Cinema Prices in Italy. In P. Del Giovane and R. Sabbatini (Eds.), *The Euro, inflation and Consumers' Perceptions. Lessons from Italy*, pp. 125–157. Berlin-Heidelberg, Germany: Springer.

Chakroun, M. (2010). Health Care Expenditure and GDP: An International Panel Smooth Transition Approach. *International Journal of Economics 4*(1), 189–200.

Choi, I. (2001). Unit Root Tests for Panel Data. *Journal for International Money and Finance 20*(2), 249–272.

Colletaz, G. and C. Hurlin (2006). Threshold Effects in the Public Capital Productivity: An International Panel Smooth Transition Approach. *Document de Recherche LEO 2006-04*.

Curtin, R. (2007). What U.S. Consumers Know About Economic Conditions. *University of Michigan Working Paper*.

Del Giovane, P., S. Fabiani, and R. Sabbatini (2008). What's Behind "Inflation Perceptions"? – A Survey-Based Analysis of Italian Consumers. *Banca d'Italia Eurosistema Working Paper 655*.

Destais, G., J. Fouquau, and C. Hurlin (2009). Energy Demand Models: A Threshold Panel Specification of the "Kuznets Curve". *Applied Economic Letters 16*, 1241–1244.

Döhring, B. and A. Mordonu (2007). What Drives Inflation Perceptions? A Dynamic Panel Data Analysis. *European Economy Economic Papers 284*.

Döpke, J., J. Dovern, U. Fritsche, and J. Slacalek (2008). The Dynamics of European Inflation Expectations. *B. E. Journal of Macroeconomics 8*(1), Article 12.

Dräger, L. (2011). Inflation Perceptions and Expectations in Sweden – Are Media Reports the "Missing Link"? *KOF Swiss Economic Institute Working Paper 273*.

Dufour, J. (1981). Rank Tests for Serial Dependence. *Journal of Time Series Analysis 2*(3), 117–128.

Dziuda, W. and G. Mastrobuoni (2009). The Euro Changeover and Its Effects on Price Transparency and Inflation. *Journal of Money, Credit and Banking 41*(1), 101–129.

Ehrmann, M. (2006). Rational Inattention, Inflation Developments and Perceptions after the Euro Cash Changeover. *ECB Working Paper Series 588*.

Eife, T. A. (2006). *Currency Changeovers as Natural Experiments*. Ph. D. thesis, Ludwig-Maximilians-Universität München.

Eife, T. A. and W. T. Coombs (2007). Coping with People's Inflation Perceptions During a Currency Changeover. *University of Heidelberg – Department of Economics – Discussion Paper Series 458*.

Engle, R. F. and C. W. J. Granger (1987). Co-integration and error-correction: Representation, estimation, and testing. *Econometrica 55*(2), 251–276.

Fehr, E. and J.-R. Tyran (2007). Money Illusion and Coordination Failure. *Games and Economic Behavior 58*(2), 246–268.

Fisher, I. (1928). *The Money Illusion*. New York: Adelphi Company.

Foellmi, R., R. Rosenblatt-Wisch, and K. Schenk-Hoppe (2011). Consumption Paths Under Prospect Utility in an Optimal Growth Model. *Journal of Economic Dynamics & Control 35*(3), 273–281.

Fok, D., D. van Dijk, and P. H. Franses (2005). A Multi-Level Panel STAR Model for US Manufacturing Sectors. *Journal of Applied Econometrics 20*(6), 811–827.

Fouquau, J., C. Hurlin, and I. Rabaud (2008). The Feldstein-Horioka Puzzle: a Panel Smooth Transition Regression Approach. *Economic Modelling 25*(2), 284–299.

Fullone, F., M. Gamba, E. Giovannini, and M. Malgarini (2007). What Do Citizens Know about Statistics? The Results of an OECD/ ISAE Survey on Italian Consumers. *OECD/ISAE Manuscript*.

Gaffeo, E., I. Petrella, D. Pfajfar, and E. Santoro (2011). Loss Aversion and the Transmission of Monetary Policy. *Working Paper*.

Gonzalez, A., T. Teräsvirta, and D. van Dick (2005). Panel Smooth Transition Regression Models. *SSE/EFI Working Paper Series in Economics and Finance 604*.

Granger, C. W. J. (1969). Prediction with a Generalized Cost Function. *Operations Research 20*, 199–207.

Granger, C. W. J. and T. Teräsvirta (1993). *Modelling Nonlinear Economic Relationships*. Oxford University Press.

Gutierrez, L. (2003). On the Power of Panel Cointegration Tests: a Monte Carlo Comparison. *Economics Letters 80*(1), 105–111.

Hansen, B. E. (1999). Threshold Effects in Non-Dynamic Panels: Estimation, Testing, and Inference. *Journal of Econometrics 93*(2), 345–368.

Hardie, B. G. S., E. J. Johnson, and P. S. Fader (1993). Modeling Loss Aversion and Reference Dependence Effects on Brand Choice. *Marketing Science 12*(4), 378–394.

Hoffmann, J., H.-A. Leifer, and A. Lorenz (2006). Index of Perceived Inflation or EU Consumer Surveys? An Assessment of Professor H. W. Brachinger's Approach. *Intereconomics 41*(3), 142–150.

Ibarra, R. and D. Trupkin (2011). The Relationship between Inflation and Growth: A Panel Smooth Transition Regression Approach. *Mimeo*.

Im, K. S., M. H. Pesaran, and Y. Shin (2003). Testing for Unit Roots in Heterogeneous Panels. *Journal of Econometrics 115*(1), 53–74.

Jonung, L. (1981). Perceived and Expected Rates of Inflation in Sweden. *American Economic Review 71*(5), 961–968.

Jonung, L. and C. Conflitti (2008). Is the Euro Advantageous? Does it Foster European Feelings? Europeans on the Euro after Five Years. *European Economy Economic Papers 313*.

Jonung, L. and D. Laidler (1988). Are Perceptions of Inflation Rational? Some Evidence for Sweden. *American Economic Review 78*(5), 1080–1087.

Jude, E. (2010). Financial Development and Growth: A Panel Smooth Regression Approach. *Journal of Economic Development 35*, 15–33.

Jungermann, H., H. W. Brachinger, J. Belting, K. Grinberg, and E. Zacharias (2007). The Euro Changeover and the Factors Influencing Perceived Inflation. *Journal of Consumer Policy 30*(4), 405–419.

Kahneman, D. and A. Tversky (1979). Prospect Theory: An Analysis of Decision under Risk. *Econometrica 47*(2), 263–291.

Kao, C. (1999). Spurious Regression and Residual-Based Tests for Cointegration in Panel Data. *Journal of Econometrics 90*(1), 1–44.

Lamla, M. J. and S. M. Lein (2008). The Role of Media for Consumers' Inflation Expectation Formation. *KOF Swiss Economic Institute Working Paper 201.*

Lein, S. M. and T. Maag (2011). The Formation of Inflation Perceptions: Some Empirical Facts for European Countries. *Scottish Journal of Political Economy 58*(2), 155–188.

Levin, A., C.-F. Lin, and C.-S. J. Chu (2002). Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties. *Journal of Econometrics 108*(1), 1–24.

Maag, T. (2009). On the Accuracy of the Probability Method for Quantifying Beliefs about Inflation. *KOF Swiss Economic Institute Working Paper 230.*

Maddala, G. S. and S. Wu (1999). A Comparative Study of Unit Root Tests with Panel Data and a New Simple Test. *Oxford Bulletin of Economics and Statistics 61*(Special Issue), 631–652.

Malgarini, M. (2008). Quantitative Inflation Perceptions and Expectations of Italian Consumers. *ISAE Working Paper 90.*

Malmendier, U. and S. Nagel (2011). Depression Babies: Do Macroeconomic Experiences Affect Risk Taking? *The Quarterly Journal of Economics 126*(1), 373–416.

Mehrara, M., M. Musai, and H. Amiri (2010). The Relationship between Health Expenditure and GDP in OECD Countries Using PSTR. *European Journal of Economics, Finance and Administrative Sciences 24*, 50–58.

Mincer, J. and V. Zarnowitz (1969). The Evaluation of Economic Forecasts. In J. Mincer (Ed.), *Economic Forecasts and Expectations: Analysis of Forecasting Behavior and Performance*, pp. 1 – 46. NBER: NBER.

Muth, J. F. (1961). Rational Expectations and the Theory of Price Movements. *Econometrica 29*(3), 315–335.

Nardo, M. (2003). The Quantification of Qualitative Survey Data: A Critical Assessment. *Journal of Economic Surveys 17*(5), 645–668.

Nickell, S. (1981). Biases in Dynamic Models with Fixed Effects. *Econometrica 49*(6), 1417–1426.

Nielsen, H. (2003). Inflation Expectations in the EU - Results from Survey Data. *Humboldt-Universität zu Berlin, Sonderforschungsbereich 373: Quantification and Simulation of Economic Processes 13.*

Patton, A. J. and A. Timmermann (2007). Testing Forecast Optimality Under Unknown Loss. *Journal of the American Statistical Association 102*, 1172–1184.

Pedroni, P. (1999). Critical Values for Cointegration Tests in Heterogenous Panels With Multiple Regressors. *Oxford Bulletin of Economics and Statistics 61*(Special Issue), 653–670.

Pedroni, P. (2001). Purchasing Power Parity Tests in Cointegrated Panels. *The Review of Economic and Statistics 83*(4), 727–731.

Pedroni, P. (2004). Panel Cointegration: Asymptotic and Finite Sample Properties of Pooled Time Series Tests With an Application to the PPP Hypothesis. *Econometric Theory 20*(3), 597–625.

Pesaran, M. H. (2004). General Diagnostic Tests for Cross Section Dependence in Panels. *Cambridge Working Paper in Economics 0435*.

Pesaran, M. H. (2007). A Simple Panel Unit Root Test in the Presence of Cross-Section Dependence. *Journal of Applied Economics 22*(2), 265–312.

Rosenblatt-Wisch, R. (2008). Loss Aversion in Aggregate Macroeconomic Time Series. *European Economic Review 52*(7), 1140–1159.

Rotemberg, J. J. (2005). Customer Anger at Price Increases, Changes in the Frequency of Price Adjustment and Monetary Policy. *Journal of Monetary Economics 52*(4), 829–852.

Rotemberg, J. J. (2008). Behavioral Aspects of Price Setting, and Their Policy Implications. *NBER Working Paper 13754*.

Shafir, E., P. Diamond, and A. Tversky (2004). Money Illusion. In C. Camerer, G. Loewenstein, and M. Rabin (Eds.), *Advances in Behavioral Economics*, Chapter 17, pp. 483–509. Princeton, New Jersey: Princeton University Press.

Starmer, C. (2004). Developments in Nonexpected-Utility Theory: The Hunt for a Descriptive Theory of Choice under Risk. In C. Camerer, G. Loewenstein, and M. Rabin (Eds.), *Advances in Behavioral Economics*, Chapter 4, pp. 104–147. Princeton, New Jersey: Princeton University Press.

Thaler, R. (1985). Mental Accounting and Consumer Choice. *Marketing Science 4*(3), 199–214.

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*(6), 739–760.

Tversky, A. and D. Kahneman (1981). The Framing of Decisions and the Psychology of Choice. *Science 211*(4481), 453–458.

Tversky, A. and D. Kahneman (1991). Loss Aversion in Riskless Choice: A Reference-Dependent Model. *Quarterly Journal of Economics 106*(4), 1039–1061.

van der Klaauw, W., W. Bruine de Bruin, G. Topa, S. Potter, and M. Bryan (2008). Rethinking the Measurement of Household Inflation Expectations: Preliminary Findings. *Federal Reserve Bank of New York Staff Reports 359*.

Wilcoxon, F. (1945). Individual comparisons by ranking methods. *Biometrics Bulletin 1*(6), 80–83.

Tables

Table 1: Quantile Test for Rationality of Inflation Perceptions

	1996m1–2010m12		1996m1–2001m12		2002m1–2010m12		1996m1–2007m12	
Country	z-stat	Prob.	z-stat	Prob.	z-stat	Prob.	z-stat	Prob.
AT	3.973	***	-3.120	***	0.24		3.662	***
BE	3.027	***	0.351		0.19		5.259	***
ES	3.274	***	2.099	**	0.98		5.790	***
FI	-0.450		na	na	-2.28	**	0.634	
FR	3.153	***	-2.276	**	0.24		3.884	***
GER	2.978	***	1.151		2.16	**	2.951	***
GR	3.976	***	3.729	***	4.54	***	3.464	***
IT	4.430	***	3.001	***	3.83	***	3.517	***
NL	7.401	***	-1.673	*	5.27	***	7.238	***
PT	4.798	***	1.996	*	4.04	***	6.822	***

***, **, * denote significance at the 1%, 5% and 10% level, respectively.

Table 2: PSTR–Model: Linearity Tests

No of thresholds m	π^p_{t-1} linear (1)	(2)	π^p_{t-1} nonlinear (1)	(2)
	transition variable: inflation			
$H_0 : r = 0$ vs. $H_1 : r = 1$	0.082	7.729	2.044	13.224
	(0.775)	(0.000)	(0.130)	(0.000)
$H_0 : r = 1$ vs. $H_1 : r = 2$	-	-	9.280	10.459
	-	-	(0.000)	(0.000)
$H_0 : r = 2$ vs. $H_1 : r = 3$	-	-	0.121	0.002
	-	-	(0.886)	(1.000)
	transition variable: time			
$H_0 : r = 0$ vs. $H_1 : r = 1$	30.344	29.574	15.513	22.318
	(0.000)	(0.000)	(0.000)	(0.000)

Note: F-Statistic used for linearity tests. Numbers in parentheses denote p-values.
Sample period: 1996m1–2010m12.

Table 3: PSTR–Model: Determination of the Number of Thresholds

No of thresholds m	π^p_{t-1} linear (1,1)	(1,2)	(2,2)	π^p_{t-1} nonlinear (1,1)	(1,2)	(2,2)
No of param	8	9	10	10	11	12
RSS	59.971	57.678	57.948	59.241	55.961	55.926
AIC	-3.387	-3.425*	-3.419	-3.397*	-3.453	-3.453
Schwarz	-3.363	-3.397*	-3.389	-3.367*	-3.419	-3.416

Note: Test statistic used: F. Numbers in parentheses denote p-values. (i,j) defines the number of thresholds for each transition function, e.g. (1,2) estimates one threshold for inflation, and two thresholds for time. Optimal model highlighted by *.
Sample period: 1996m1–2010m12.

Table 4: PSTR Estimation Results

π_t^p	π_{t-1}^p linear (1,1)	(1,2)	π_{t-1}^p nonlinear (1,1)	(1,2)
γ_G	2.272***	1.900***	0.820	1.262
	(0.000)	(0.000)	(3.138)	(2.292)
γ_T	3066.7***	8143.0	2162.2***	9585.1
	(0.358)	(65113.2)	(181.7)	(300019.7)
$c_{1,G}$	2.668***	3.307**	1.821	3.313*
	(0.737)	(1.278)	(10.951)	(1.828)
$c_{1,T}$	0.402***	0.402***	0.403***	0.403***
	(0.004)	(0.003)	(0.005)	(0.017)
$c_{2,T}$	-	0.835***	-	0.835***
		(0.006)		(0.005)
π_{t-1}^p	0.916***	0.909***	0.827***	0.935***
	(0.005)	(0.005)	(0.020)	(0.008)
$\pi_{t-1}^p G$	-	-	0.082***	0.016
			(0.023)	(0.016)
$\pi_{t-1}^p T$	-	-	0.048***	-0.066***
			(0.011)	(0.009)
π_t	0.032***	0.092***	0.037**	0.034**
	(0.009)	(0.008)	(0.017)	(0.011)
$\pi_t G$	0.038***	0.040***	0.038**	0.060***
	(0.008)	(0.007)	(0.019)	(0.015)
$\pi_t T$	0.036***	-0.051***	0.000	0.007
	(0.004)	(0.004)	(0.001)	(0.009)

Note: Standard errors in parentheses.

***,**,* denote significance at the 1%, 5%, and 10% level.

Sample period: 1996m1–2010m12.

Table 5: Dynamic Fixed Effects Estimates of Loss Aversion

	1996m1–2010m12			1996m1–2007m12	
π_t^p	(1)	(2)	(3)	(2)	(3)
π_{t-1}^p	0.935***	0.938***	0.940***	0.957***	0.964***
	(0.007)	(0.008)	(0.008)	(0.008)	(0.008)
$\text{thold}_{1,2}$	-	-0.035	-0.067***	-0.032	-0.061***
		(0.022)	(0.023)	(0.023)	(0.023)
π_t	0.079***	0.067***	0.062***	0.048***	0.036***
	(0.007)	(0.010)	(0.011)	(0.012)	(0.013)
π_t thold_1	-	0.021*	-	0.020*	-
		(0.011)		(0.012)	
π_t thold_2	-	-	0.035***	-	0.039***
			(0.012)		(0.012)
constant	-0.051**	-0.039*	-0.037	-0.041	-0.038
	(0.022)	(0.023)	(0.023)	(0.028)	(0.028)
R^2	0.983	0.983	0.984	0.988	0.988
AIC	-889.848	-894.265	-905.318	-1088.192	-1103.791
BIC	-878.868	-872.306	-883.358	-1067.130	-1082.730
CADF test residuals	-5.686	-5.710	-5.664	-5.108	-5.064
prob.	0.000	0.000	0.000	0.000	0.000

Note: Correlated panels corrected standard errors in parentheses.
***, **, * denote significance at the 1%, 5% and 10% level, respectively.

Figures

Figure 1: Prospect Theory and Inflation Perceptions

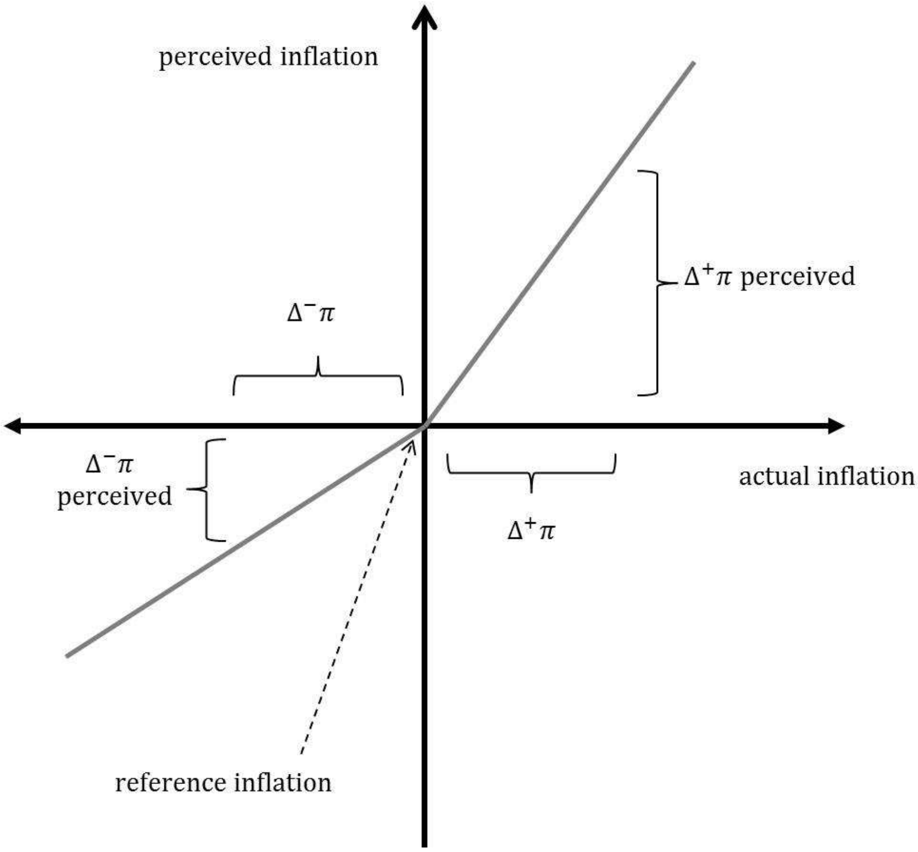


Figure 2: Actual and Perceived Inflation for all Countries

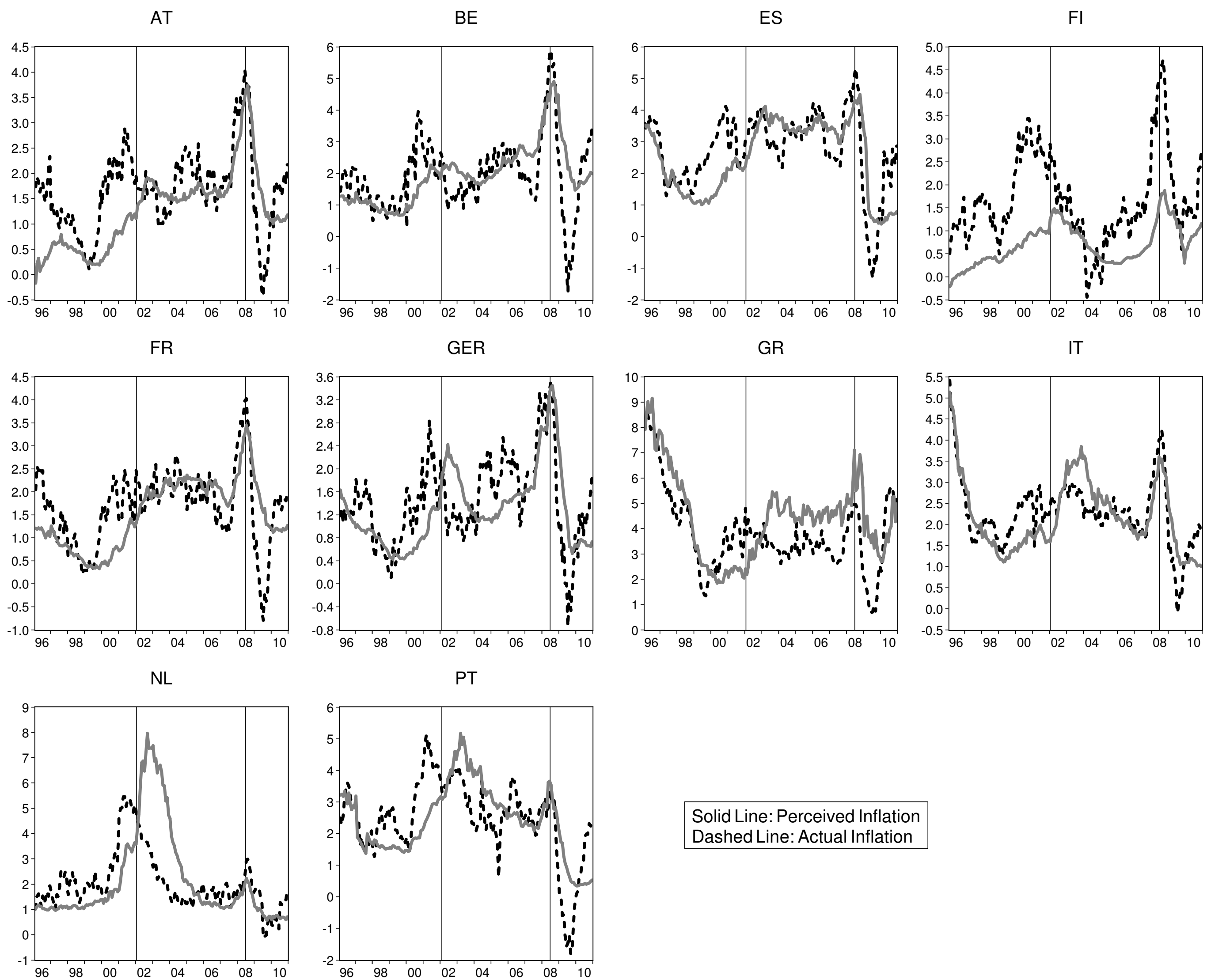


Figure 3: Inflation vs. Elasticities

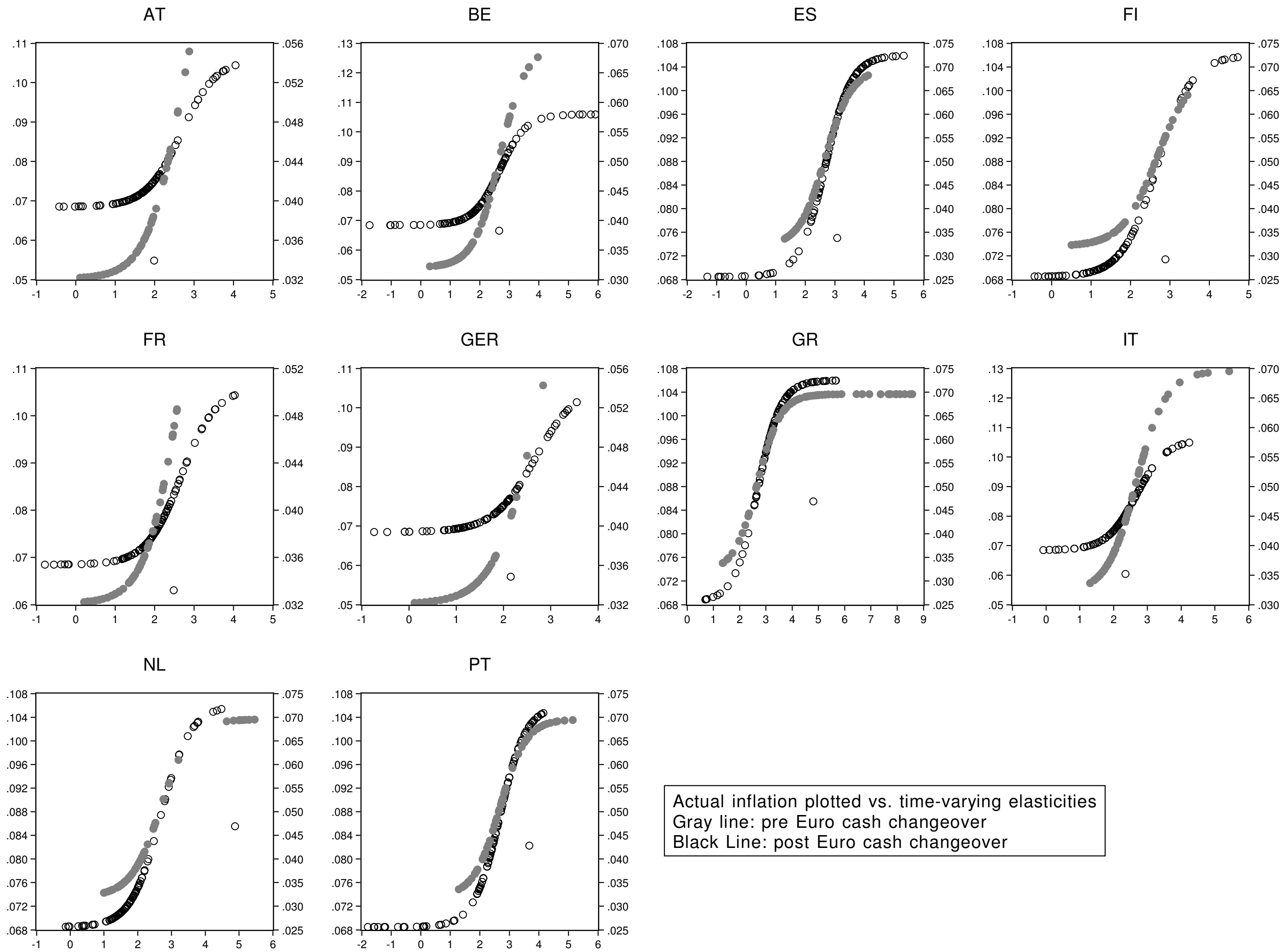
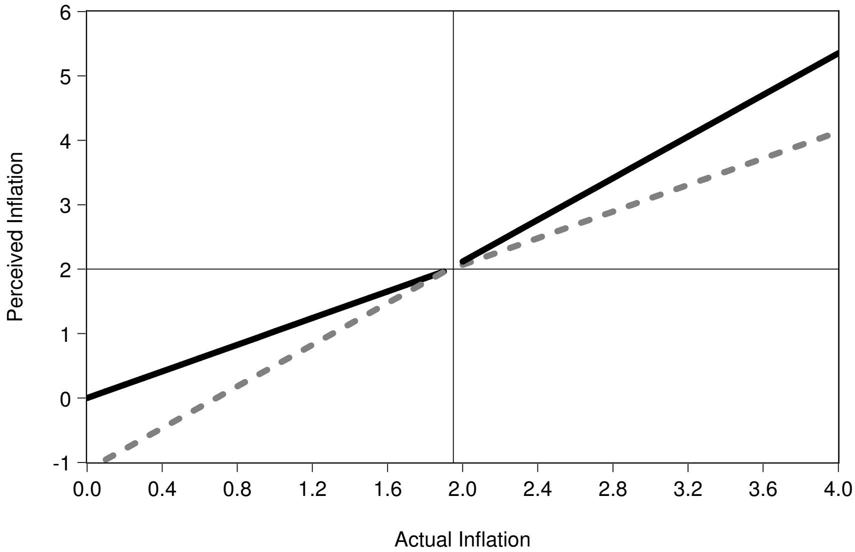


Figure 4: Estimated Value Function under Loss Aversion



Appendix

A Quantification technique

Question 5 of the EC *Joint Harmonized EU Program of Business and Consumer Surveys* asks if consumer prices over the past 12 months did:

1. Fall,
2. Stay about the same,
3. Increase at a slower rate,
4. Increase at the same rate,
5. Increase more rapidly.

In an influential paper, Batchelor and Orr (1988) derive how to transform responses from a pentachotomous survey into a measure of inflation expectations/ inflation perceptions. The method is based on and extends the seminal work of Carlson and Parkin (1975).
We follow Nielsen (2003) – see Figure 5 below – in the description of the method and use her terminology. The main assumption underlying such a method lies in the existence of an interval $\left(-\delta_t^L, \delta_t^U\right)$ around 0 with $\delta_t^L, \delta_t^U > 0$ such that participants in the survey report "no change" in prices (i.e. zero inflation). Furthermore, there exists another interval $\left(\tilde{\mu}_t - \varepsilon_t^L, \tilde{\mu}_t + \varepsilon_t^L\right)$ around the subjective expected value of the inflation rate $\tilde{\mu}_t$ such that individuals report that "prices increase at the same rate".
The respective questions of the survey here can be translated into such a concept in the following way (x_{t+1} measures the time series of interest – here: "inflation perceptions" which households have to form expectations about):[27]

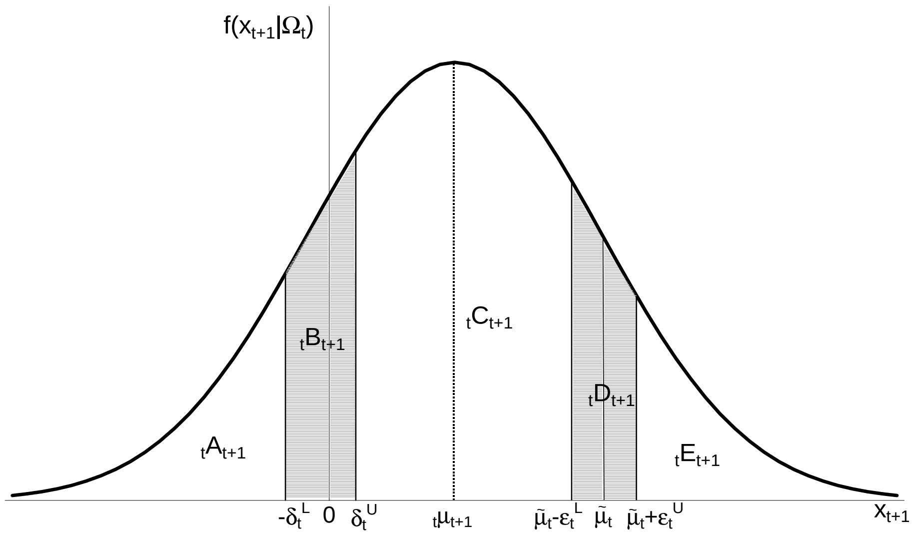
1. "Fall slightly" if $x_{t+1} \leq -\delta_t^L$.
2. "Stay about the same" if $-\delta_t^L < x_{t+1} \leq \delta_t^U$,
3. "Increase at a slower rate" if $\delta_t^U < x_{t+1} \leq \tilde{\mu}_t - \varepsilon_t^L$,
4. "Increase at the same rate" if $\tilde{\mu}_t - \varepsilon_t^L < x_{t+1} < \tilde{\mu}_t + \varepsilon_t^U$,
5. "Increase more rapidly" if $\tilde{\mu}_t + \varepsilon_t^U \leq x_{t+1}$.

[27] Note that we deal with perceptions as a latent variable which has to be "forecasted" based on the information set in t. This is in line with models of sticky information as well as models of rational inattention where updating of information is costly.

We label the fractions of answers in the ordering of the questions from A to E, respectively, that is: ${}_t\mu_{t+1} = \tilde{\mu}_t \times f\left({}_tA_{t+1}, \ldots, {}_tE_{t+1}\right)$, where ${}_tA_{t+1}, \ldots, {}_tE_{t+1}$ are the fractions of respondents answering each option, f is a known distribution function (see Batchelor and Orr, 1988, p. 322, formula (11)) and $\tilde{\mu}_t$ is the expected value of the perceived inflation rate that has to be specified.

We use a version of the procedure proposed in Döpke et al. (2008) in order to determine f and $\tilde{\mu}_t$: We assume a normal distribution for f and furthermore base $\tilde{\mu}_t$ on the medium-term trend of inflation, which households are assumed to observe correctly. This is approximated by the HP filter, which is calculated in a recursive way following a quasi-real-time approach. For each period, t, we apply the filter with the usual penalty parameter ($\lambda_{HP} = 14400$) for monthly data. Finally, we set $\tilde{\mu}_t$ equal to the value of the HP filtered inflation as of time t.

Figure 5: Quantification of a Pentachotomous Survey



Source: Nielsen, 2003, p.5

B Unit Root Tests and Cointegration

Both inflation perceptions and actual inflation rates in our panel are tested for unit roots over the whole sample period from January 1996 to December 2010. We apply five different panel unit root tests: The Levin et al. (2002) test assumes a common unit root process over all series in the sample. It estimates proxies for Δy_{it} and y_{it-1} and tests for the null hypothesis H_0 : $\alpha = 1$ in the regression $\Delta y^*_{it} = \alpha y^*_{it-1} + \eta_{it}$, allowing for individual-specific deterministic intercepts. However, it suffers from the restriction that no cross-sectional correlation is allowed and that it can only test for stationarity of all series in the sample. By contrast, the tests by Im et al. (2003), Maddala and Wu (1999) as well as Choi (2001) (Fisher's ADF and PP test) allow for individual unit root processes. They specify individual unit root tests and derive test statistics to test the null hypothesis H_0 : $\alpha_i = 0$, for all i against the alternative that at least one $\alpha_i \neq 0$. While the tests may include individual-specific short-run dynamics and deterministics such as time trends for each panel member, cross-sectional correlation between countries is still not fully accounted for. This may be a relevant issue for actual and perceived inflation rates in a panel of closely related countries, such as the European countries analyzed here. Therefore, we additionally test for panel unit roots with the Pesaran (2007) Cross-Sectionally Augmented Dickey-Fuller (CADF) test. The test computes a t-bar statistic averaging t-statistic values for $H_0 : \alpha_i = 0$ from a standard ADF-regression augmented with lagged and first-differenced values of the cross-sectional mean of the series. All panel unit root tests are calculated with three lags.

The test results in Table B.1 uniformly reject the null of a unit root in inflation in favor of the alternative of stationarity of at least some series in the panel. However, the null cannot be rejected with the stricter alternative of stationarity of all series in the Levin et al. (2002) test. Regarding the stationarity of perceived inflation rates in the panel, test results are less conclusive. The Im et al. (2003) and the Maddala and Wu (1999) tests reject the null of a unit root at the 5% level, favoring the alternative of stationarity of some series in the panel. By contrast, the Choi (2001) and, more strongly, the Pesaran (2007) CADF test cannot reject the null of a unit root. Our results are in line with findings in Lein and Maag (2011), who also find that inflation perceptions are more persistent in a similar panel setting for a shorter sample period. Generally, empirical evidence on the order of integration of inflation series is mixed, Altissimo et al. (2006) conclude in a survey that empirical findings seem to lean towards stationarity of inflation.

Due to the inconclusive evidence on stationarity of perceived and actual inflation in our panel, we furthermore test for panel cointegration between perceived and actual inflation. Results are presented for seven panel coin-

tegration test statistics proposed by Pedroni (1999, 2001, 2004) that are calculated by extending the Engle-Granger-framework to the panel setting and testing for stationarity of the residual from a regression with I(1) variables, while allowing for individual fixed effects and time trends. The null hypothesis of no cointegration ($\rho_i = 1$) is tested either against the alternative of a common cointegrating vector ($\rho_i = \rho < 1$) or against the alternative of individual cointegrating relationships ($\rho_i < 1$). The Kao (1999) panel cointegration test is also residual-based, but does not allow for individual-specific deterministics. Stationarity of the residuals from the first-stage regression is then tested with a panel ADF test on the null of no cointegration against the alternative of a common cointegrating vector. Finally, the Maddala and Wu (1999) test computes individual Johansen cointegration trace tests and maximum eigenvalue tests and uses those to obtain a combined Fisher statistic. Gutierrez (2003) conducts a Monte Carlo experiment to compare the power of Kao (1999) and Pedroni (1999, 2001, 2004) tests and finds that as T gets large, the Pedroni tests have higher power than the Kao test.
Results shown in Table B.2 give strong evidence of a cointegration relation between perceived and actual inflation over the whole sample period. All Pedroni tests, as well as the Kao (1999) test, reject the null of no cointegration in favor of the alternative of either a common or individual cointegrating vectors at the 1% level. While the Maddala and Wu (1999) trace test and maximum-eigenvalue test rejects both the null of no and the null of at most one cointegration relation, test statistics for the latter are significantly smaller. Our, quite intuitive, result of cointegration between perceived and actual inflation is in line with findings in Lein and Maag (2011) who also report panel cointegration between perceptions and inflation with a slightly different sample.

Table B.1: Panel Unit Root Tests for Perceived and Actual Inflation

	Perceptions		Inflation	
Method	Stat.	Prob.*	Stat.	Prob.*
Alternative: Stationarity of all series in the panel				
Levin, Lin & Chu t	-1.496	0.067	-0.098	0.461
Alternative: Stationarity of some series in the panel				
Im, Pesaran and Shin W-stat	-2.184	0.015	-4.185	0.000
ADF - Fisher Chi-square	33.742	0.028	53.513	0.000
PP - Fisher Chi-square	27.530	0.121	66.159	0.000
Pesaran CADF t-bar	-1.986	0.245	-2.594	0.002

* Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality.
Sample period: 1996m1–2010m12.

Table B.2: Panel Cointegration Tests between Perceived and Actual Inflation

Method	1996m1–2010m12	
Pedroni Tests	(Weighted) Stat.	Prob.
Alternative hypothesis: common AR coefs.		
Panel v-Statistic	9.015	0.000
Panel rho-Statistic	-6.627	0.000
Panel PP-Statistic	-4.270	0.000
Panel ADF-Statistic	-3.571	0.000
Alternative hypothesis: individual AR coefs.		
Group rho-Statistic	-4.957	0.000
Group PP-Statistic	-4.058	0.000
Group ADF-Statistic	-3.270	0.001
Kao ADF Test	t-Stat.	Prob.
	-6.154	0.000
Maddala & Wu Test	Fisher-Stat.*	Prob.
Trace test		
None	202.600	0.000
At most 1	89.980	0.000
Max.-Eigenvalue test		
None	165.500	0.000
At most 1	89.980	0.000

* Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. Chi-square distribution. All other tests assume asymptotic normality.

C Rationality Tests under Symmetric Loss

We applied three types of rationality tests:

- A traditional Mincer-Zarnowitz regression where the perception error is regressed on a constant and lagged perception errors.
- As a test for bias, we applied the non-parametric test of Dufour (1981) and Campbell and Ghysels (1995) which has the advantage of avoiding restrictive assumptions on well-behaved residuals as in the case of most of the regression-based tests. Suppose that we compute the absolute value of the difference between each observation and the mean, and then rank these observations from high to low. Then the test is based on the idea that the sum of the ranks for the samples above and below the median should be similar. We can use a Wilcoxon signed rank test with the null hypothesis that the median of the perception errors is equal to zero. According to Campbell and Ghysels (1995), this test is preferable to parametric tests for small samples.
- As a test for efficiency, we again use a test proposed by Dufour (1981) and Campbell and Ghysels (1995) to test for absence of serial correlation in the forecasts errors, i. e. that the median of the product of two consecutive perception errors is centered on a median of zero.

We do not report the Mincer-Zarnowitz results here as the result is well-known from other studies, see for instance Jonung and Laidler (1988), Lein and Maag (2011) and Dräger (2011). Rationality under symmetric loss is rejected in all cases, irrespective of controlling for possible structural breaks at the end of 2001 or 2007. The results, however, are available from the authors upon request.
This rejection is probably to a large extent driven by the "inefficiency channel". The results of the two non-parametric tests shown in table (C.1) revealed that in a number of cases we cannot reject unbiasedness once we control for possible structural breaks by splitting the sample appropriately. However, efficiency is rejected in all cases irrespective of sample splits.

Table C.1: Nonparametric Rationality Tests: Unbiasedness and Efficiency

Unbiasedness

Country	1996m1–2010m12 Stat.	Prob.	1996m1–2001m12 Stat.	Prob.	2002m1–2010m12 Stat.	Prob.	1996m1–2007m12 Stat.	Prob.
AT	6.840	***	7.287	***	1.884	*	8.040	***
BE	0.921		6.114	***	3.321	***	1.892	*
ES	3.585	***	6.602	***	0.887		3.924	***
FI	10.827	***	7.371	***	7.557	***	9.542	***
FR	3.522	***	6.467	***	1.053		4.231	***
GER	3.219	***	4.756	***	0.379		4.492	***
GR	6.153	***	1.086		6.533	***	5.362	***
IT	2.098	**	4.823	***	1.203		1.987	**
NL	2.746	***	7.354	***	2.224	**	1.638	
PT	0.542		5.530	***	3.664	***	1.630	

Efficiency

Country	1996m1–2010m12 Stat.	Prob.	1996m1–2001m12 Stat.	Prob.	2002m1–2010m12 Stat.	Prob.	1996m1–2007m12 Stat.	Prob.
AT	10.608	***	7.165	***	7.554	***	9.499	***
BE	10.153	***	6.185	***	8.095	***	8.993	***
ES	10.690	***	6.902	***	8.089	***	9.450	***
FI	11.418	***	7.320	***	8.589	***	10.222	***
FR	9.944	***	6.804	***	7.277	***	8.304	***
GER	10.569	***	6.627	***	8.225	***	9.557	***
GR	11.307	***	6.907	***	8.887	***	10.081	***
IT	10.823	***	6.844	***	8.284	***	9.573	***
NL	11.240	***	7.251	***	8.524	***	10.236	***
PT	10.478	***	6.294	***	8.381	***	9.353	***

Test statistic: Wilcoxon signed rank.

***, **, * denote significance at the 1%, 5% and 10% level, respectively.

D Structural Break Tests

Table D.1: Quandt-Likelihood-Ratio Test for Structural Breaks

Country	Lossaversion with thold1			Lossaversion with thold2		
	Max. Wald F	p-value	date	Max. Wald F	p-value	date
AT	53.892	0.000	2008m09	53.008	0.000	2008m09
BE	30.449	0.000	2008m06	33.299	0.000	2008m06
ES	16.170*	0.055	2003m07	25.271*	0.002	2006m06
FI	23.190	0.008	2008m04	24.821	0.004	2007m12
FR	30.074	0.000	2002m02	37.286	0.000	2008m03
GER	42.598	0.000	2007m08	41.840	0.000	2007m09
GR	23.495	0.007	2008m07	18.962**	0.011	2002m09
IT	22.528	0.010	2008m07	23.272	0.007	2008m07
NL	25.205	0.003	2002m10	36.786	0.000	2002m10
PT	13.157	0.247	1998m06	9.152*	0.451	2001m06

Note: *, ** denotes 30%, 40% trimmed data, respectively. The default is 15% trimmed data.
Sample period: 1996m1–2010m12.