State-dependence vs. Time-dependence: An Empirical Multi-Country Investigation of Price Sluggishness

by Steffen Ahrens and Matthias Hartmann

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Keywords: State- vs. Time-Dependence, Phillips Curve, Functional Coefficients.

JEL classification: E31, E58, C14.

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Preliminary Version

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1 Introduction

The effectiveness of monetary policy to impact real variables such as employment and output crucially depends on the extent to which prices react sluggishly to central banks' policy innovations. Therefore, mechanisms of price sluggishness have become a central aspect of modern DSGE models. In terms of the standard NKM, price sluggishness strongly influences aggregate inflation dynamics yielding a non-vertical NKPC in the short run.

The most widely used price-updating mechanism is the Calvo (1983) staggered contracts model, where a constant, randomly selected fraction of firms adjust their prices at each time instance in a monopolistically competitive market. Despite its popularity, this time-dependent specification has been frequently criticized as being a rather restrictive description of the price setting process (Caplin and Leahy, 1991; Wolman, 1999). In particular, indicative evidence by Fernández-Villaverde and Rubio-Ramírez (2010) suggests that the Calvo parameter should not be regarded as a structural parameter in the sense of a "deep" and state-invariant coefficient.

The contribution of this paper is an empirical investigation of the behavior of aggregate price setting. We implement a testing procedure by means of a nonparametric representation of the structural form NKPC. Such a functional-coefficient regression model allows to express the Calvo parameter as a functional coefficient which may be systematically affected by observable factor variables such as inflation, inflation uncertainty (IU, henceforth), or both factors simultaneously (Danziger, 1983). This specification nests both the typically employed

time- and state-dependent pricing rules. This corresponds to testing for the null hypothesis of parameter constancy (time-dependent pricing) against the alternative hypothesis of inflation- or IU-induced price revisions.

For this purpose, we obtain a so-called functional coefficient representation of the NKPC. This semiparametric model class allows to express functional dependence of parameters on observable factor variables (Cai et al., 2000). An important advantage of our approach is that it allows to draw inference on the state-dependence of the pricing scheme by taking potential heteroscedasticity of the disturbances into account, which is particularly critical in models which relate price adjustment to inflation or IU (Sims, 2001). To address residual heteroscedasticity in the inference process regarding functional dependence, we make use of the so-called factor-based bootstrap (Herwartz and Xu, 2009).

The distinction between time- and state-dependent pricing schemes is of crucial importance from the policy maker's point of view. The welfare implications, measured by minimizing an objective function which is quadratic in inflation and the output gap (Woodford, 2003), under both schemes generally do not coincide (Lombardo and Vestin, 2008). Applying the Calvo model to a state-dependent world, monetary policy runs the risk of putting too little weight on inflation stabilization.

To summarize the most important findings, we first confirm assertions frequently made in theoretical discussions that the Calvo scheme is a rather restrictive model of aggregate price setting. Moreover, it is documented that a number of shortcomings of empirical NKPC model representations in explaining inflation data may be addressed by means of a state-dependent pricing rule. In particular, problems of insignificant or even implausibly negative estimates of the relation between inflation and marginal costs are considerably reduced in the framework

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of our more general NKPC specification. The remainder of this paper is organized as follows. Section 2 reviews the relevant literature. Section 3 describes the data set and introduces the model framework. Subsequently, our approach to estimation and inference is introduced. Section 4 summarizes and discusses the empirical results. Section 5 concludes.

2 Relation to the Literature

Mechanisms of price sluggishness can be assigned to either time-dependent models of price setting or state-dependent models of price setting. In time-dependent price setting models, firms change prices in discrete (Taylor, 1979) or random (Calvo, 1983) time intervals, independent of the underlying economic environment. In contrast, state-dependent price setting models assume price adjustments to be somewhat costly[1] and therefore price changes depend on observable fundamental economic factors such as inflation or IU (Fabiani et al., 2006). Although, especially the newer state-dependent models (e.g. Golosov and Lucas, 2007; Gertler and Leahy, 2008; Costain and Narkov, 2011a,b; Dotsey et al., 2009; Midrigan, 2011) reasonably well resemble a fair amount of the stylized facts of price setting behavior, time-dependent models - especially the Calvo (1983) model - are still the most widely adopted price updating schemes in the literature on monetary policy. The straightforward reason is their analytical elegance and tractability.

For time-dependent price setting models firms' price changing decisions are independent of economic fundamentals - including inflation and IU -, but depends exclusively on time.

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[1]These costs can take a variety of different forms, e.g. physical adjustment or "menu" costs (Sheshinski and Weiss, 1977; Rotemberg, 1982; Mankiw, 1985; Golosov and Lucas, 2007; Gertler and Leahy, 2008; among many others), information costs (Reis, 2006; Maćkowiak and Wiederholt, 2009; Woodford, 2009), or consumer costs such as customer disenchantment (Sibly, 2002, 2007), customer anger (Rotemberg, 2005), and customer regret (Rotemberg, 2010).

This is in stark contrast with theoretical and empirical evidence. The theoretical literature suggests inflation to have a positive influence on the frequency of price adjustment (Sheshinski and Weiss, 1977[2]; Naish, 1986; Ball et al., 1988; Romer, 1990; Golosov and Lucas, 2007). In the context of DSGE models, Fernández-Villaverde and Rubio-Ramírez (2010) find that movements in the Calvo pricing parameter are negatively correlated with inflation. Canova (2006) estimates a small-scale KM for a variety of data samples for the United States and reports that the Calvo parameter seems to be relatively stable over most subsamples, with some variation for a few subsamples. Also Cogley and Sbordone (2005) find some weak evidence for variation of the Calvo parameter over different time periods for a non-zero steady state NKPC with indexation and strategic complementarities. These theoretical predictions are strongly supported by empirical evidence of, e.g. Klenow and Kryvtsov (2008), Nakamura and Steinsson (2008), and Klenow and Malin (2010) for the United States and Álvarez et al. (2006), Dyhne et al. (2006), and Vermeulen (2012) for the euro area. For IU the picture is less clear cut. While IU might be used to cover increases in firms' markup (Van Hoomissen, 1988; Bénabou, 1992; Tomassi, 1994) and thereby increases the frequency of price adjustments, it might also induce stronger search effort by customers, which leads to a closer monitoring of prices and consequently reduces price changes (Bénabou, 1992; Bénabou and Konieczny, 1994). Furthermore, a negative impact of IU on the frequency of price adjustment is also apparent in the presence of price adjustments costs (Sheshinksi and Weiss, 1983; Danziger, 1999).

The distinction between time- and state-dependent pricing schemes is of crucial impor-

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[2]Sheshinski and Weiss (1977) show that in general inflation has an ambiguous effect on the frequency of price changes. The negative effect of inflation on price changes occurs, however, only under unreasonably high inflation rates.

tance from the policy maker's point of view. First, the output effects of monetary shocks are typically stronger and longer lasting for time-dependent models relative to state-dependent models (Dotsey et al., 1999; Golosov and Lucas, 2007; Gertler and Leahy, 2008; Midrigan, 2011). Second, the welfare implications, measured by minimizing an objective function which is quadratic in inflation and the output gap (Woodford, 2003), under both schemes generally do not coincide (Lombardo and Vestin, 2008).[3]

Our empirical approach is based on the semiparametric estimation of a so-called functional coefficient model. This allows to express functional dependence of parameters on observable factor variables (Cai et al., 2000). This method enables us to test for state-dependence of the Calvo parameter. An important advantage of this approach is that we can draw inference on the state-dependence of the pricing scheme by taking potential heteroscedasticity of the disturbances into account. This is required in models which relate price adjustment to inflation or IU, since such processes are characterized by conditional heteroscedasticity (Sims, 2001). If prices are more flexible at higher inflation rates or IU, this is likely reflected in the conditional volatility of inflation (Cogley and Sargent, 2005; Fernández-Villaverde and Rubio-Ramírez, 2007). To address residual heteroscedasticity in the inference process regarding functional dependence, we make use of the recently proposed factor-based bootstrap (Herwartz and Xu, 2009). This scheme resamples factor observations in contrast to drawing from the residuals as it is common, e.g. in the typically employed residual bootstrap. We describe the bootstrap scheme in detail after the introduction of the estimation method.

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[3]This holds even true for the comparison of Calvo (1983) and Rotemberg (1982) pricing, which up to a first-order approximation around the zero inflation steady state, result in observationally equivalent reduced-form macroeconomic dynamics (Roberts, 1995).

3 Empirical Approach

3.1 Data

The data set comprises quarterly observations of real output, the implicit output deflator, and unit labor costs for $N = 14$ advanced economies, namely Australia, Belgium, Canada, Finland, France, Italy, Japan, Netherlands, New Zealand, Portugal, Spain, Sweden, the United Kingdom, and the United States, from 1961Q3 to 2011Q4 taken from the OECD Economic Outlook No. 90. All series are seasonally adjusted. Inflation is defined as quarterly percentage change, i.e. $\pi_t = 400 \times (p_t - p_{t-1})$ with p_t denoting the natural logarithm of the implicit output deflator and $t = 1, ..., T$ representing the time instances between 1961Q3 and 2011Q4, i.e. $T = 201$. Since real marginal costs mc_t are unobservable, we follow the suggestion of Galí and Gertler (1999) and use the labor's share of income s_t as proxy instead, i.e. $mc_t = s_t$. The labor's share of income is equivalent to real unit labor costs and in log-linearized terms given by $s_t = ulc_t - p_t$, with ulc_t being nominal unit labor cost. Finally, we follow the mainstream procedure in the macroeconomic literature on estimating NKPC and generate the output gap $\tilde{y}_t = y_t - \bar{y}_t$ by applying the Hodrick-Prescott filter with smoothing parameter $\lambda = 1600$ to the series of real output y_t, which obtains the long-run trend estimate $\bar{y}_t$ (Galí and Gertler, 1999; Galí et al., 2001).

3.2 Model Framework

Recent microeconometric studies on pricing behavior show that neither time-dependent nor state-dependent models alone are capable of fully replicating the various patterns of price movements in the data (Klenow and Kryvtsov, 2008; Nakamura and Steinsson, 2008; Klenow

and Malin, 2010). Nevertheless, the majority of monetary policy analysis is conducted in NKMs resting on purely time-dependent pricing mechanisms such as the prominent Calvo (1983) staggered pricing scheme (e.g. Clarida et al., 1999; Eggertsson and Woodford, 2003; Smets and Wouters, 2003; Schmitt-Grohe and Uribe, 2007; among others). According to the Calvo scheme, each period individual firms have a certain probability $(1 - \theta)$ to be allowed to reset their price, while with probability θ they have to remain their previous price. In the aggregate, such pricing behavior leads to the New Keynesian recitation of the Phillips curve, which relates inflation to expected future inflation and a measure of real marginal costs

$$\check{\pi}_t = \beta E_t \check{\pi}_{t+1} + \frac{(1 - \theta)(1 - \theta\beta)}{\theta} \check{mc}_t, \tag{1}$$

where $\check{\pi}_t$ denotes inflation, $\check{mc}_t$ represents real marginal costs and $\beta < 1$ is a discount factor. Moreover, the Calvo probability $\theta \in [0, 1]$ determines the degree of price inertia, where $\theta = 0, 1$ refers to cases of fully flexible and fully rigid prices, respectively. In the model of Calvo (1983), the average duration of non-adjustment amounts to a fixed spell of $1/(1 - \theta)$ quarters for the aggregate price level. Substitution of the expectation error $\varepsilon_t = \beta[E_t[\check{\pi}_{t+1}] - \check{\pi}_{t+1}]$ under rational expectation yields

$$\check{\pi}_t = \beta \check{\pi}_{t+1} + \frac{(1 - \theta)(1 - \theta\beta)}{\theta} \check{s}_t + \varepsilon_t, \tag{2}$$

In the framework of the NKPC, both, the specification in equation (2) and the shorthand representation, which is obtained by letting $\kappa \equiv ((1 - \theta)(1 - \theta\beta))/\theta$, have an economic interpretation. Galí and Gertler (1999) refer to κ as a "reduced form" parameter and distinguish this quantity from the "structural" coefficients of the NKPC from equation (2). As the term

"structural" indicates, the price adjustment speed parameter θ is treated as a constant, i.e. θ is assumed to be independent of any economic fundamentals. However, allowing the frequency of price adjustment to co-vary with economic fundamentals influences the reduced form parameter κ and thus leads to a change in the sensitivity of inflation to innovations in real marginal cost (Gertler and Leahy, 2008) and hence to changes in the central banks ability to stabilize inflation via the nominal interest rate.

To allow for such non-constant behavior of the Calvo parameter, we employ a state-dependent NKPC, where the frequency of price adjustment depends on economic fundamentals rather than solely on time. The result is that the Calvo parameter $\theta(\omega)$ is a function of ω, where ω represents potential factors variables. The simplest way to introduce such state-dependence into the Calvo (1983) mechanism is to allow firms to choose their optimal stochastic arrival rate θ, given a cost of changing price. Such an approach has been introduced, among others, by Romer (1990), Kiley (2000), Devereux and Yetman (2002), and Levin and Yun (2007). In this context the authors derive a state-dependent Calvo parameter $\theta(\pi)$, with $\frac{\partial \theta(\pi)}{\partial(\pi)} < 0$ (Bakhshi et al., 2007b).[4] Bakhshi et al. (2007a) show that the Calvo purely time-dependent NKPC, equation (1), is a special case of a more general Calvo state-dependent NKPC with $\theta(\omega)$. Therefore, we apply the generalization $\theta(\omega)$ to equation (1) and refer to equation (3) as our state-dependent NKPC. Thus, equation (3) reads

$$\check{\pi}_t = \beta \check{\pi}_{t+1} + \frac{(1 - \theta(\omega))(1 - \theta(\omega)\beta)}{\theta(\omega)} \check{s}_t + \varepsilon_t, \tag{3}$$

where $\omega = (w^{(1)}, w^{(2)})$, i.e. we allow for bivariate state-dependence of the Calvo parameter.

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[4]A related widespread approach to derive a state-dependent NKPC based on the Calvo mechanism is presented by Dotsey et al. (1999).

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This formulation may be employed to detect changes in firms' price setting behavior which are driven by potential factor variables $w_t^{(\bullet)}$, where '$\bullet = 1, 2$' indicates (1) lagged inflation π_{t-1} and (2) lagged inflation uncertainty IU_{t-1}. Inflation uncertainty is defined as $IU_{t-1} = |\Delta \pi_{t-1}| = |\pi_{t-1} - \pi_{t-2}|$, i.e. the absolute error of the inflation forecast from a random walk model. Such predictions are frequently found to obtain superior predictive performance as compared to other inflation forecasting schemes (Canova, 2007; Stock and Watson, 2007, 2008). To account for different scales of the inflation and IU processes, $w_t^{(\bullet)}$ is considered in standardized form, i.e. $w_t^{(\bullet)} = \widetilde{w}_t^{(\bullet)}/\sigma(\widetilde{w})$ with $\sigma(\widetilde{w})$ denoting the standard error of $\widetilde{w}_t^{(\bullet)}$.

To examine the potential factor dependence of the Calvo parameter θ, the influence of $\breve{\pi}_{t+1}$ on $\breve{\pi}_t$ and $\breve{mc}_t$ is accounted for by means of a partial regression step prior to the introduction of the state-dependent NKPC. To isolate the effect of $\breve{\pi}_{t+1}$ on $\breve{mc}_t$, we let $\boldsymbol{\breve{mc}} = (\breve{mc}_1, ..., \breve{mc}_T)'$, $\boldsymbol{\breve{\pi}} = (\breve{\pi}_1, ..., \breve{\pi}_T)'$, and $\boldsymbol{\breve{\pi}}_+ = (\breve{\pi}_2, ..., \breve{\pi}_{T+1})'$, assuming that one additional observation is available. Then, $\boldsymbol{mc} = (I_T - \boldsymbol{\breve{\pi}}_+(\boldsymbol{\breve{\pi}}'_+\boldsymbol{\breve{\pi}}_+)^{-1}\boldsymbol{\breve{\pi}}'_+)\boldsymbol{\breve{mc}}$ where I_T denotes the identity matrix of dimension T, whereas $\boldsymbol{\pi} = \boldsymbol{\breve{\pi}} - \beta\boldsymbol{\breve{\pi}}_+$ may be obtained by presetting $\beta = 0.99$. Such magnitudes of the discount parameter β are commonly calibrated for quarterly data (Smets and Wouters, 2003; Altig et al., 2005; Sbordone, 2005; Dufour et al., 2006). Estimation of β also yields values close to 0.99 (Galí and Gertler, 1999; Dufour et al., 2006). Accounting for the effect of π_{t+1} in this way results in an equivalent representation of equation (3). The condensed representation is advantageous since we focus on the state-dependence of θ. The state-dependent NKPC is given by

$$\pi_t = \frac{(1 - \theta(\omega))(1 - \beta\theta(\omega))}{\theta(\omega)}mc_t + e_t, \tag{4}$$

where e_t denotes the error term in the regression after controlling for the effect of $\tilde{\pi}_{t+1}$ on $\check{pi}_t$ and $\check{mc}_t$.

3.3 Estimation

Estimation of the factor dependent price adjustment frequency proceeds in analogy to the semiparametric Nadaraya Watson estimation method (Nadaraya, 1964; Watson, 1964). Thereby, we express functional dependence of the price adjustment parameter on π_{t-1} and IU_{t-1}. Apart from potential state-dependence, the employed estimation procedure has to take account of the potential endogeneity of mc_t, which is standard practice in the related literature, where estimation of the NKPC is discussed (see Galí and Gertler, 1999; Sbordone, 2005; and the references therein). The estimation of the NKPC commonly proceeds by means of the generalized method of moments (GMM). In the framework of the functional coefficient model (4), we account for regressor endogeneity by estimating $\theta(\omega)$ according to

$$\hat{\theta}(\omega) = \arg\min_{\theta} q(\theta, K_h(\omega)), \tag{5}$$

with $q(\cdot)$ denoting the GMM objective function

$$q(\theta, K_h, \omega) = \bar{m}(\cdot)'\Phi\bar{m}(\cdot), \tag{6}$$

where $K_h(u) = K(u/h)/h$, with $K(\cdot)$ being a kernel function depending on the so-called bandwidth parameter $h > 0$. Moreover, Φ represents a positive definite weighting matrix and

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$\bar{m}(\cdot)$ is shorthand for the (empirical) moment condition

$$\bar{m}(\theta, K_h, \omega) = (1/T) \sum_{t=1}^{T} z_t e_t K_h(w_t^{(1)} - w^{(1)}) K_h(w_t^{(2)} - w^{(2)}). \tag{7}$$

In equation (7), z_t represents a vector of instrument variables.

3.4 Implementation

Theoretical descriptions of how price adjustment responds to π or IU suggest that nominal rigidity is decreasing for higher inflation rates and in cases of rising IU (Ball et al., 1988). If the response of θ to π or IU is not excessively volatile, observations $w_t^{(\bullet)}$ near point $w^{(\bullet)}$ should be informative for the value of the functional $\theta(w^{(1)}, w^{(2)})$ near $w^{(\bullet)}$ (Eubank, 1988; Härdle, 1990). The closer observations $w_t^{(\bullet)}$ are to a point $w^{(\bullet)}$, the more informative they will typically be regarding the behavior of the functional $\theta(\omega)$ near $w^{(\bullet)}$. These differences in the predictive content are incorporated in the estimation by means of the kernel function, which puts higher relative weight on those observations in proximity to $w^{(\bullet)}$. In equation (4), the relation between π_t and mc_t is evaluated in a neighborhood of ω by means of the kernel weighting function $K_h(\cdot)$. Estimation of $\hat{\theta}(\omega)$ yields local averages of the hypothesized state-dependent relation. An important part of semiparametric regression is the choice of h. This parameter determines how the tradeoff between unbiasedness and efficiency of estimation is addressed. While smaller bandwidths tend to increase the variability of estimates, larger values may hide local characteristics of the relation between θ and ω. For increasing h, $\hat{\theta}(\omega)$ approaches the limit of the usual time-invariant GMM estimate. This highlights that the functional coefficient method is suitable to contrast systematic variation in θ from time invariance,

since the NKPC under the latter assumption is nested in the state-dependent regression model (4). We choose the bandwidth according to Scott's rule of thumb (Scott, 1992), which obtains as $h = 1.06 T^{-1/5}$, since the factor variables are considered in standardized form. We employ the logistic Kernel, i.e. $K(u) = \Lambda(u)/(1 - \Lambda(u))$, where $\Lambda(u) = 1/(1 + \exp(-u))$. For the graphical display of the functional dependence, $\theta(w^{(1)} = v^{(1)}, w^{(2)} = v^{(2)})$ is evaluated at particular states $(v^{(1)}, v^{(2)})$ from the equidistant grid

$$v^{(\bullet)} = c_{\text{lo}}^{(\bullet)}, c_{\text{lo}}^{(\bullet)} + k\mathcal{L}^{(\bullet)}, ..., c_{\text{up}}^{(\bullet)}, \; k = 1, 2, ... \, , \tag{8}$$

where $c_{\text{lo}}^{(\bullet)}, c_{\text{up}}^{(\bullet)}$ denote lower and the upper quantiles of the factor observations $w_t^{(\bullet)}$, $t = 1, ..., T$ and $\mathcal{L}^{(\bullet)}$ determines the step length. Particular choices of quantiles from $w_t^{(\bullet)}$, $t = 1, ..., T$, are determined to facilitate the graphical exposition and numerical accuracy of results. Functional coefficient estimates feature highest local efficiency at the center of a (unimodal) empirical factor distribution. In our case, the sample period covers observations from higher inflation regimes from the more distant past. Corresponding levels of π have only in few instances been observed during recent times. A choice of $\{c_{\text{lo}}, c_{\text{up}}\} = \{0.2, 0.8\}$ determines a range of inflation and IU which is currently observed in most advanced economies. This can be seen from Figure 1, where estimates of the empirical density function of inflation are depicted. In the left plot, density estimates for π_{it}, $i = 1, ..., 14$, $t = 1, ..., T$ are shown. The plot on the right shows respective kernel estimates for the IU series. Dashed lines indicate the cutoff points, which are determined as $\{c_{\text{lo}}, c_{\text{up}}\} = \{0.01, 0.8\}$ as a suitable range of IU for which local dependence of θ is examined.
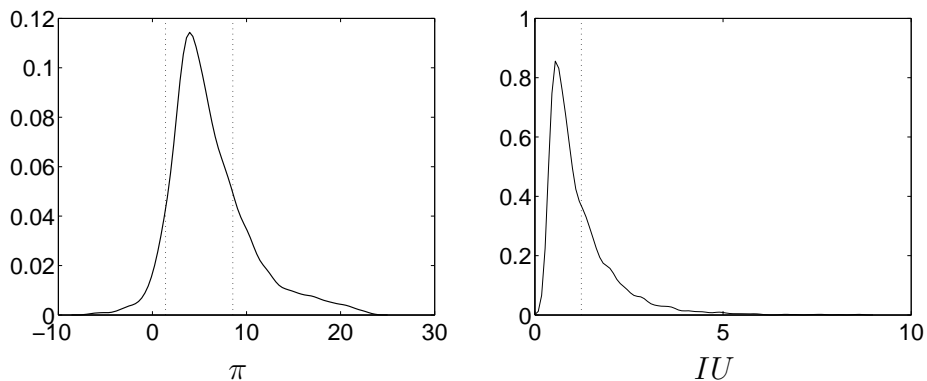
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Figure 1: Inflation rates and IU: Smoothed empirical densities of factor observations

3.5 Inference

In the framework of the functional coefficient NKPC, we intend to test if the adjustment parameter is constant or state-dependent. In the literature on functional coefficient estimation, such tests are routinely implemented by means of bootstrap approaches, i.e. by resampling from the disturbance term (Cai et al., 2000). The conclusions drawn from this resampling scheme, however, might be affected by heteroscedasticity in the disturbances (Herwartz and Xu, 2009). This is particularly relevant, since changes in the variance of inflation series over time are empirically well documented for a wide range of economies (Engle, 1982; Hartmann and Herwartz, 2012). For this reason, we employ the so-called factor-based bootstrap as suggested by Herwartz and Xu (2009), which is designed to circumvent the problems encountered by residual-based resampling procedures in case of heteroscedastic disturbances.

1. Functional coefficients evaluated at particular realizations of the data and for a given choice of h may be described as

$$\hat{\theta}(\omega) = \theta \left(\pi_t, mc_t, \omega_t = (w_t^{(1)}, w_t^{(2)}), h, t = 1, ..., T \right).$$ (9)

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2. To distinguish state-dependence from structural constancy in the pricing scheme, local estimates $\hat{\theta}(\omega)$ are compared to their bootstrap counterparts

$$\hat{\theta}^*(\omega) = \theta\left(\pi_t, mc_t, \omega_t^* = (w_t^{(1*)}, w_t^{(2*)}), h, t = 1, ..., T\right), \tag{10}$$

with binary tuples $(w_t^{(1*)}, w_t^{(2*)})$ being drawn with replacement from the factor observations $(w_t^{(1)}, w_t^{(2)})$.

3. A large number as, e.g. $\mathfrak{R} = 1000$ resampling estimates $\hat{\theta}^*(\omega)$ obtains the bootstrap distribution of $\hat{\theta}^*(\omega)$. The corresponding confidence interval is employed to assess the local state-dependence of $\theta(\omega)$. In this study, we reject state-invariance at the 10 percent level, if $\hat{\theta}(\omega)$ is either below the 5 percent or above the 95 percent-quantile of the bootstrap distribution at any level of the factor variables.

As it can be seen from step number 2 as described above, in this approach, the bootstrap confidence intervals are obtained by imposing H_0 "directly" during the bootstrap, i.e. we distort the relation between θ and the factor observations in ω and thereby guarantee that $H_0 : \theta(\omega) = \theta \forall \omega$ holds irrespectively of potential heterogeneity in the errors (or the factor observations).

4 Results

In the following, we report estimates and test outcomes for the state-dependence of θ and we comment on the magnitudes and economic plausibility of implied estimates of the NKPC relation. Results obtained by means of pooled panel estimation, where observations for all

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economies are jointly considered are also reported. In the literature it is well documented that the considered economies feature distinctive characteristics, particularly with respect to different levels of inflation or IU (Judson and Orphanides, 1999; Caporale and Kontonikas, 2009). Therefore, the conventional pooled estimation framework might be regarded as rather restrictive. However, the functional coefficient representation captures individual economies' idiosyncratic characteristics through the influence of factor variables. This introduces considerable flexibility also in the pooled estimation setting.

Figure 2 and Figure 3 show estimates obtained according to equation (5) for the United States and the pooled sample. Solid lines represent the estimates $\hat{\theta}$, dashed lines stand for 90 percent bootstrap confidence intervals. The latter are obtained according to the factor-based bootstrap as described in Section 3.5. Local state-dependence at particular factor levels is indicated if estimates are outside the interval. For clarity, we present only a subset of estimates from the entire range of the factor space. Dependence of θ on one of the factors is plotted conditional on a certain level of the respective other factor. For example, $\theta = \theta(\pi | IU = c_{\text{up}})$ means that potentially inflation-induced variation in θ is depicted for an IU level equal to the upper quantile of the IU series.

The estimates $\hat{\theta}$ in Figure 2 reinstate the theoretical prediction that $\frac{\partial \theta(\omega)}{\partial \pi < 0}$, i.e. the frequency of price adjustment increases for higher inflation rates. This finding is also in line with recent evidence from Nakamura and Steinsson (2008) and Klenow and Kryvtsov (2008). In contrast, we do not find evidence for a uniform sign of the IU impact. This is in line with the discussion in Bénabou (1992), where both signs are described as plausible. As a robustness check, we also obtain estimates of Calvo parameters based on data for the remaining 13 single economies. For 10 out of 13 economies, an impact of either π or IU on θ is detected. Only
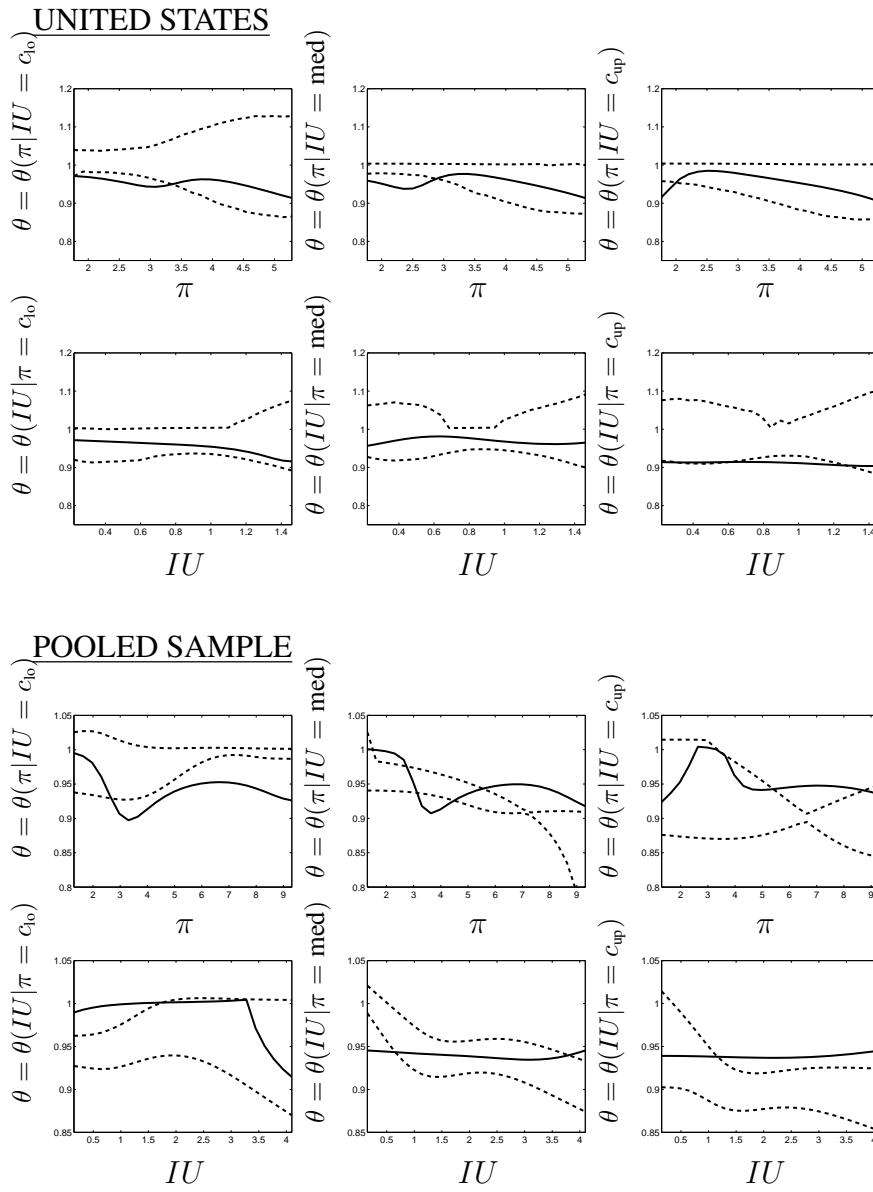
UNITED STATES

POOLED SAMPLE

Figure 2: Functional coefficient estimates for the United States and the pooled sample
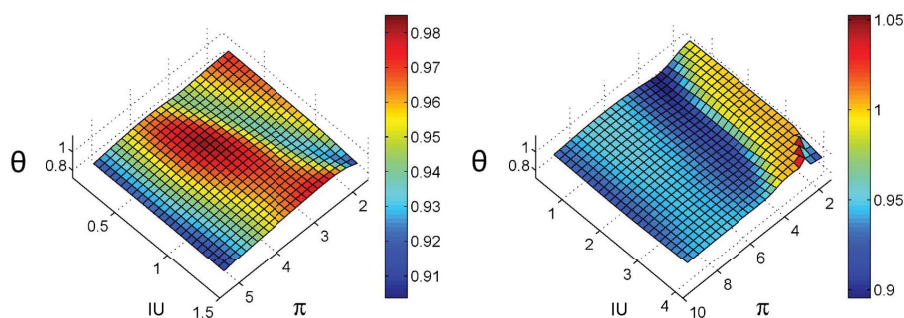
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Figure 3: Surface plots for the United States (left) and the pooled sample (right)

for Canada, Italy and Finland, the H_0 of a constant Calvo pricing scheme cannot be rejected.

In Figure 3, surface plots for the United States and the pooled estimate are depicted to provide an impression on the joint impact of π and IU on θ. Surfaces for the remaining individual economies are qualitatively similar and not reported to economize on space. Both plots of Figure 3 show that while θ takes an initially high level for low inflation rates, the estimates drop at intermediate levels of π around 3 percent. In case of the pooled estimate, the updating frequency is less responsive for much higher π. At first, the price inertia for values of π which are currently observed in most advanced economies might appear relatively high. This is in contrast to micro-price studies which find averagely fixed prices between one and two quarters (Bils and Klenow, 2004; Klenow and Kryvtsov, 2008; Nakamura and Steinsson, 2008; Klenow and Malin, 2010). One reason for this divergence is the use of different observational frequencies. While the above mentioned studies use monthly consumer price index time series in their estimations, this paper applies quarterly aggregates of the GDP deflator. Ellis (2009) and Abe and Tonogi (2010) show that lower frequency data leads to larger estimates of price stickiness by construction. Micro-price studies at very high frequencies, such as weekly or even daily, report price spells of less than a quarter (Kehoe and Midrigan, 2007; Ellis, 2009;

	ARCH(1)	ARCH(4)	$J \times 10^3$		ARCH(1)	ARCH(4)	$J \times 10^3$
AT	41.70	52.07	0.09	JP	18.89	26.01	0.01
BE	9.24	24.41	0.01	NL	32.37	111.74	0.32
CA	28.61	33.40	0.03	NZ	41.31	53.47	7.21
ES	34.94	42.81	0.01	PT	38.44	52.26	0.04
FN	74.89	102.17	0.97	SW	76.26	81.09	0.01
FR	42.67	51.31	0.51	UK	77.99	79.46	0.00
IT	26.25	42.10	0.12	US	43.99	51.42	0.01

Table 1: Regression diagnostics

Abe and Tonogi, 2010; Cavallo, 2012). Moreover, estimating a high-frequency NKPC, Ahrens and Sacht (2014) show that also on the macro level higher-frequency data leads to lower average price spells.

The magnitude of θ, however, is close to estimates reported in other studies which investigate aggregate pricing (Smets and Wouters, 2003; Levin et al., 2006; Eichenbaum and Fisher, 2007; Nason and Smith, 2008; Fernández-Villaverde and Rubio-Ramírez, 2010). The influence of IU on θ is in both cases confined to moderate inflation rates. However, this range of inflation is also currently most frequently observed. Whereas higher IU leads to decreasing θ in the United States for low π, the effect is ambiguous in case of the pooled estimate. This suggests that IU influences θ in a rather idiosyncratic way.

In Table 1, diagnostic test statistics are summarized. These statistics are obtained for estimates of equation (2) assuming no state-dependence of θ. Columns 1-2 and 4-5 report ARCH-LM test statistics (Engle, 1982) for the residuals from estimation of equation (2) with $q = 1, 4$ denoting the lag order of squared disturbances. These ARCH-LM tests confirm the presence of conditional heteroscedasticity in the residuals for each considered economy. Our findings are in line with the findings of Fernández-Villaverde and Rubio-Ramírez (2010), who point out that ARCH-effects might lead to spurious conclusions regarding state-dependence

or dynamics in θ. Similarly, residual-based bootstrap methods as considered by, e.g., Cai et al. (2000) are unreliable in cases when disturbances feature ARCH dynamics (Herwartz and Xu, 2009). In such a situation, the factor-based bootstrap approach might be a more suitable means to draw inference on functional dependence of coefficients. Furthermore, Columns 3 and 6 report J-t est statistics for overidentifying restrictions in the GMM estimation procedure. The J-statistics in Table 1 indicate no evidence against the null hypothesis of joint exogeneity of the instrument variable (IV) set. We choose $z_t = (\tilde{y}_{t-1}, \tilde{y}_{t-2})'$ as instrument variables, a subset of the instrument variables considered by, e.g. Galí and Gertler (1999), where $\tilde{y}_t = y_t - \bar{y}_t$ denotes the output gap, i.e. the deviation of gross domestic product y_t from its long term trend $\bar{y}_t$.

With 2 instrument variables, the J-test for overidentification adheres to a $\chi^2(1)$ distribution under H_0 of at least one of the instrument variables being exogenous. Depending on initial examination of the J-statistic, we determine the IV set alternatively as $z_t = \tilde{y}_{t-1}$ in cases where exogeneity is rejected.

A further way to assess the plausibility of the obtained estimates is to examine the magnitude and significance of the reduced-form parameter $\kappa \equiv ((1-\theta)(1-\theta\beta))/\theta$. A puzzling finding of many studies, where similar to Galí and Gertler (1999) the labor's share of income is employed as an explanatory variable in the structural NKPC is that estimates of κ are insignificant or even have a theoretically implausible negative sign (Jondeau and Bihan, 2005; Rudd and Whelan, 2005; Abbas and Sgro, 2011; Kuttner and Robinson, 2012).

Table 2 shows reduced form Phillips curve estimates κ and corresponding t-statistics based on Newey and West (1987) standard errors, as they are typically reported in related studies. For the economies we consider, the sign of the Phillips curve relation is positive, as predicted

	κ	t-stat.		κ	t-stat.
AT	0.01	0.01	JP	0.01	0.01
BE	0.02	0.02	NL	0.01	0.01
CA	0.01	0.01	NZ	0.02	0.11
ES	0.02	0.03	PT	0.02	0.25
FN	0.01	0.02	SW	0.03	0.02
FR	0.02	0.03	UK	0.05	0.06
IT	0.01	0.11	US	0.01	0.01

Table 2: Estimates for equation (2) (constant θ case)

by economic theory. The magnitudes of estimates are for all economies similar to the findings reported by Galí and Gertler (1999), Galí et al. (2001), or Sbordone (2005), among many others. Moreover, in line with existing empirical evidence, none of the coefficients is statistically significant. Since disturbances are found to be heteroscedastic, t-statistics are based on a robust covariance estimator (Newey and West, 1987).

The recurring finding of implausible NKPC parameter estimates has led to doubts about the suitability of the labor's share of income as a measure of marginal costs (Wolman, 1999; Neiss and Nelson, 2002; Kiley, 2007). The criticism put forth in these studies is also based on theoretical arguments. However, Wolman (1999) and Galí et al. (2005) point out that it might be the overly restrictive assumption of a constant price updating frequency, as implied by the Calvo (1983) scheme, that gives rise to estimation problems. This hypothesis can be addressed by means of the functional coefficient framework. In analogy to the investigation described above, we estimate the reduced-form NKPC, allowing for state-dependence such that $\kappa = \kappa(\omega)$. Since functional dependence of θ is detected in the majority of economies, the same might also hold for κ.[5] Local estimates of κ and corresponding t-statistics for distinct

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[5]Functional coefficient estimates which allow for state-dependence of both θ and the discount parameter β suggest that β is not affected by either π or IU. These results are not reported in detail and might be obtained from the authors upon request.
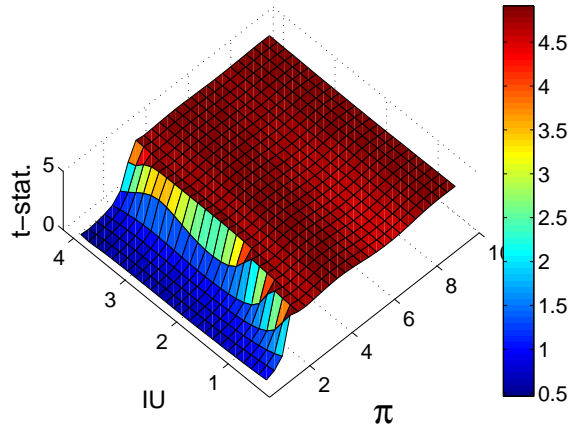
Figure 4: State-dependent t-statistics for the reduced-form Phillips curve parameter κ

levels of π and IU indicate if the generalization reinstates the theory with empirical NKPC estimates. We find that allowing for state-dependence of κ obtains estimates at similar magnitudes as reported in Table 2. The t-statistics are mostly higher than their counterparts in Table 2 but are, however, throughout insignificant also in this case. However, insufficient degrees of freedom might deteriorate the power of t-tests regarding local semiparametric estimates to a larger extent than in the parametric case. We, therefore, compare pooled estimates under the assumption of a constant and state-dependent κ. As depicted in Figure 4, the t-statistics for functional coefficient estimates of κ are highly significant over almost the entire range of the factor space. The respective state-invariant pooled t-statistic, in contrast, is equal to $t_{\text{pooled}} = 1.04$. Though significance tests for individual economies are not rejected, these findings are at least an indication that state-dependence is a meaningful generalization of the Calvo scheme.

For macroeconomic theory the results obtained above are particularly noteworthy, since most studies on monetary policy are conducted in a time-dependent, rather than a state-

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dependent, framework. The reason is straightforward: there seems to be a widespread agreement in the literature that both approaches are almost equivalent (Ascari and Rossi, 2012) and time-dependent models are analytically simpler and much more tractable. The equivalency result is true, however, only under very restrictive assumptions. For instance, Roberts (1995) shows that the standard approaches to time- and state-dependent pricing (which are the time-dependent approaches by Taylor (1979) and Calvo (1983) and the state-dependent approach by Rotemberg (1982)) yield observationally equivalent reduced-form dynamics up to a first-order Taylor approximation around the zero-inflation steady state. Ascari and Rossi (2012), however, show that this does not hold true anymore in the presence of trend inflation. Also, the welfare implications under these approaches, measured by minimizing an objective function which is quadratic in inflation and the output gap (Woodford, 2003), coincide up to a second-order Taylor approximation only as long as the steady state is efficient[6] (Nisticò, 2007; Lombardo and Vestin, 2008; Damjanovic and Nolan, 2011).

Finally, the output effects of monetary shocks are typically stronger and longer lasting for time-dependent models relative to state-dependent models (Dotsey et al., 1999; Golosov and Lucas, 2007; Gertler and Leahy, 2008; Midrigan, 2011). Intuitively, the reason is straightforward. While a positive monetary policy shock increases inflation, this in turn increases the price updating frequency. With more prices being updated, the Phillips curve flattens and the output reaction ceases. In time-dependent models, as noted, the average frequency of nominal adjustment is independent of inflation. Additionally, state-dependent models feature a selection effect, which is not met by time-dependent models. In state-dependent models those firms change prices, whose prices are most out of line. Therefore, nominal adjustments are

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[6]In the sense that distortions from monopolistic competition are met by a subsidy to the firm, which elevates the quantity supplied to the level under perfect competition.

quite large compared to the adjustment under time-dependent models. Consequently, state-dependent models feature a much stronger nominal flexibility (Caplin and Spulber, 1987; Golosov and Lucas, 2007).

5   Summary and Concluding Remarks

In this paper, the method of functional coefficient regression is applied to investigate on the state-dependence of the frequency of price updating. We find that both the inflation rate and IU significantly affect aggregate price adjustment. Inference is based on a bootstrap methodology which is unaffected by heteroscedasticity in the regression disturbances. Nonspherical disturbances are described as a principal impediment to valid inference in previous empirical examinations of state-dependent pricing rules. We find that the updating frequency increases at higher inflation rates. Moreover, functional coefficient estimates of the Phillips curve relation are found to be more in line with theory than estimates obtained under the assumption of constant coefficients. These finding imply that the "deep parameter" interpretation of the standard Calvo (1983) price setting scheme is a too restrictive assumption for actual price setting behavior.

Our results are of particular importance for the conduct of monetary policy analysis. First, the welfare implications under time- and state-dependent approaches coincide only under the very restrictive assumption of an efficient steady state, which is unlikely to be met. Second, output effects of monetary shocks are typically stronger and longer lasting for time-dependent models relative to state-dependent models. Therefore, the correct application of time- or state-dependent pricing schemes to the particular economy of interest is of crucial importance from

the policy maker's point of view.

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