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Conventional and unconventional monetary policy reaction to uncertainty in advanced economies: evidence from quantile regressions

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Abstract:
This paper investigates how the Federal Reserve (Fed) and the Bank of England, Bank of Japan and the European Central Bank reacted in the aftermath of the financial crisis by making use of both conditional and unconditional interest rate quantiles regressions and data on shadow short rate of interest and a measure of uncertainty. Firstly, the unconditional quantile regression offers some support for increased reaction by the Fed as the ZLB is approached. Secondly, the decreased reaction of the Fed and other monetary policy makers towards uncertainty particularly at lower conditional quantiles of interest rates lends support to expansionary mechanism in place during this time. Hence uncertainty is key to policy reaction, and more so during episodes of crisis.

Keywords: advanced economies, conditional and unconditional quantile regressions, interest rate rule, shadow rate of interest, uncertainty, zero lower bound

JEL classification: C22, E52

DOI: 10.1515/snde-2018-0056

1 Introduction

Empirical work on the analysis of monetary policy is dominated by studies that use the linear Taylor rule after Taylor (1993) and Clarida, Gali, and Gertler (2000), with prominent studies that have estimated asymmetric monetary policy reaction functions.[1] In the context of quantile regression, the response of interest rate to inflation and output (and exchange rate) has been investigated at various quantiles on the conditional distribution of interest rates. See Christou et al. (forthcoming) for a detailed discussion in this regard. This method has particular appeal, in an attempt to gauge the policy makers' reactions at the zero and lower bound amidst recent financial crisis and economic events. Chevapatrakul, Kim, and Mizen (2009) use conditional interest rate quantile regression applied to Taylor rules for Japan and the USA. They found that inflation has a larger effect on higher quantiles of interest rates than at lower quantiles of interest rates contrary to the greater aggression to inflation that they expected to find as interest rates reach low levels as the lower bound is approached. Wolters (2012) found significant and systematic variations of parameters over the conditional quantile interest rate distribution for the US. More recently, Chen and Kashiwagi (2017), using a sample that includes recent periods of zero interest rates for Japan highlight the potential downward biased of using uncensored quantile regressions versus censored ones. Finally, Liu (2018), using a Markov-switching conditional interest rate quantile distribution Taylor-rule model showed that the Fed responded more aggressively to inflation at upper tails than at lower tails in both monetary Dovish and Hawkish regimes, with it responding to inflation more aggressively during the latter regimes than the former across quantiles.[2]

Moreover, the impact of uncertainty on the Fed's monetary reaction function has received ample scrutiny both theoretically and empirically. The concept of uncertainty in monetary policy practice was coined by Brainard (1967) and hence the Brainard's attenuation principle. The former Federal Reserve Chairman, Greenspan (2003), contends that "Uncertainty is not just an important feature of the monetary policy landscape, it is the defining characteristic of that landscape." Mishkin (2010) laments the unfortunate reality that most existing studies on optimal monetary policy have abstracted from considerations of macroeconomic risk in the context of financial disruptions.

The Brainard attenuation principle hypothesizes that uncertainty dampens the monetary authorities' response to the target variables of monetary policy compared to when monetary policy decisions are made under complete certainty or certainty equivalence.[3]

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Kohn (1996), Kuttner and Posen (2001), and Bernanke and Reinhart (2004) among others have pointed to a more aggressive reaction of the central bank when interest rates approach the ZLB in order to guard against the possibility of deflation and negative demand shocks.

A point of departure from previous studies on these economies is that the empirical estimates are conducted not only at the central mean of interest rate, but we also take into account the response of interest rate to inflation, output and a measure of uncertainty at various points on both the conditional and unconditional distribution of interest rates.[4] This allows us to test predictions of nonlinearity in terms of the response of the central banks at different bounds of interest rate. Furthermore, to contribute to the literature on the response of monetary authorities at the zero or lower bounds, one should be cautious of interpreting the lower (upper) quantile of the conditional distribution of interest which does not necessarily correspond to a low (high) level of shadow interest rate. As for distinguishing the difference between the unconditional quantile regression and the conditional quantile regression, we follow Firpo, Fortin, and Lemieux (2009) for unconditional quantile regression in guiding us in making more precise statements.

To this end, we make use of the shadow interest rates, which is the nominal interest rate that would prevail in the absence of its effective lower bound, with it derived by modelling the term structure of the yield curve. Thus, we are also able to study the monetary policy behaviour across conventional and unconventional monetary regimes without worrying about explicitly modelling the ZLB, as the shadow short rates (SSR) offer an excellent description of the historic interest rate behaviour (Wu and Xia, 2016).[5] Understandably, by using the SSR and the quantiles based approach, with lower quantiles of the estimated monetary policy function econometrically capturing the reaction of central banks corresponding to the abnormal ZLB period (Liu 2018), we do not need to pursue any kind of censoring as in Chen and Kashiwagi (2017).[6]

Another main contribution is to add to the debate by assessing how the Fed, and other monetary authorities of advanced economies, have reacted to a (news-based) measure of uncertainty in the aftermath of the recent financial crisis. In the process, we, for the very first-time, contribute to the existing literature on quantile-based estimation of Taylor rules by analysing the behaviour of major central banks to not only inflation and output gap, but also to uncertainty, across the periods of conventional and unconventional monetary policy decisions. In sum, while studies like Evans et al. (2015), Caggiano, Castelnuovo, and Nodari (2018), and Ma, Olson, and Wohar (2018) have included uncertainty in the Taylor-rules, their sample have been restricted to the pre-ZLB era only and results derived based on a conditional-mean model, we use a quantiles-based approach that allows us to study the entire conditional distribution of the interest rate response. In the process, we also add to the literature on quantiles-based Taylor rule of Chevapatrakul, Kim, and Mizen (2009), Wolters (2012), Chen and Kashiwagi (2017), and Liu (2018) by incorporating the role of uncertainty in the model, which has been shown to be important in interest rate setting behaviour by Evans et al. (2015) and Ma, Olson, and Wohar (2018) in conditional-mean based monetary policy rules.

We found two main results, firstly, the unconditional quantile regression offers some support for increased reaction by the Fed as the ZLB is approached. Secondly the decreased reaction of the Fed and other monetary policy makers towards uncertainty particularly at lower conditional quantiles of interest rates lends support to expansionary mechanism in place during this time and the Brainard attenuation principle. Hence gauging the importance of uncertainty to policy reaction is key and more so during episodes of crisis.

The remainder of the paper is organized as follows: Section 2 discusses the methodology, while Section 3 presents the data, results and robustness checks, and finally Section 4 concludes.

2 Methodology: the interest rate rule estimation using quantile regression

For the purpose of this study, we extend Taylor's original rule in order to capture the reaction of monetary policy to economic policy uncertainty movement. Furthermore, we adopt the quantile regression methodology suggested by Lee (2007) to estimate the following monetary policy rules over the whole conditional distribution of interest rates:

$$i_t = \alpha_0(\tau) + \alpha_\pi(\tau)\pi_{t+k|t} + \alpha_y(\tau)y_t + \alpha_{epu}(\tau)epu_t + u_t, \tag{1}$$

where ε_t is a policy shock, y_t is the output gap, $\pi_{t+k|t}$ is a k-period-ahead inflation forecast and epu_t is the economic policy uncertainty index, and τ the quantile.[7]

2.1 Quantile regression analysis: the control function approach of Lee (2007)

We estimate the forward-type interest rate rule using the quantile regression methodology suggested by Lee (2007): IVQ. The methodology adjusts for endogeneity by adopting a control function approach. The model has the form:

$$i_t = x_t\alpha_x(\tau) + z'_{1t}\alpha_z(\tau) + u_t, \tag{2}$$

$$x_t = m(\tau^*) + z_t{}'\vartheta(\tau^*) + v_t, \tag{3}$$

$$Q_{u_t|x_t,z_t}(\tau\,|\,x_t, z_t) = Q_{u_t|v_t,z_t}(\tau|v_t, z_t) = Q_{u_t|v_t}(\tau|v_t) \equiv \lambda(v_t), \tag{4}$$

$$Q_{v_t|z_t}(\tau^*|z_t) = 0, \tag{5}$$

where i_t is the dependent variable, x_t is the endogenous explanatory variable, $z_t = (z_{1t}, z_{2t})'$ is the vector of exogenous explanatory variables, and u_t and v_t are unobserved random variables. Parameters $m(\tau^*)$ and $\vartheta(\tau^*)$ are unknown parameters, and $\alpha_x(\tau)$ and $\alpha_z(\tau)$ are unknown parameters of interest at for some $\tau \in (0, 1)$, $\tau^* \in (0, 1)$, while $\lambda(\cdot)$ is a real-valued function of v_t. Let us denote the τ^{th} quantile of A conditional on $C = c$ is denoted by $Q_{A\,|\,C}(\tau\,|\,c)$. According to (3) and (4), we get:

$$Q_{i_t|x_t,z_t}(\tau|x_t, z_t) = x_t\alpha_x(\tau) + z'_{1t}\alpha_z(\tau) + \lambda(v_t), \tag{6}$$

$$Q_{x_t|z_t}(\tau^*|z_t) = m(\tau^*) + z'_t\vartheta(\tau^*). \tag{7}$$

This suggests that the parameters of interest can be estimated in two steps. In the first step, we estimate v_t by the residuals of a linear $\tau^{*\text{th}}$ linear quantile regression of x_t on $(1, z_t)$. Specifically, $\hat{v}_t = x_t - \hat{m}(\tau^*) + z'_t\hat{\vartheta}(\tau^*), t = 1, 2, \ldots, T$, and $(\hat{m}(\tau^*), \hat{\vartheta}(\tau^*))$ is the solution to:

$$\min_{m,\vartheta} T^{-1} \sum_{t=1}^{T} \xi_{\tau^*}(x_t - m(\tau^*) + z'_t\vartheta(\tau^*)), \tag{8}$$

where $\xi_{\tau^*}(\kappa) = |\kappa| + (2\tau^* - 1)\kappa$, $\kappa \in (0, 1)$, the check function. In the second step, we estimate the parameters of interest in a quantile regression of i_t on x_t, z_{1t}, and the residuals $\hat{v}_t$ extracted in step 1. Lee (2007) carries out this step via series estimation. Specifically, $\lambda(v)$ can be approximated by a linear combination of smooth functions $\{f_i\colon i = 1, 2, \ldots\}$.[8] Let $\hat{w} = (x, z_1, \hat{v})$ and $F_n(\hat{w}) = (x_t, z_{1t}, f_1(\hat{v}_t), f_2(\hat{v}_t), \ldots., f_n(\hat{v}_t))$ for any positive integer n. Then, we calculate $\hat{\Delta} = (\hat{\alpha}_x(\tau), \hat{\alpha}'_z(\tau), \hat{\delta}'(\tau))'$ by solving the following minimisation problem:

$$\min_{\Delta} T^{-1} \sum_{t=1}^{T} h(\hat{w}_t)\xi_\tau[y_t - F'_n(\hat{w}_t)\Delta], \tag{9}$$

where $h(\hat{w}_t) = \mathbb{1}(\hat{w}_t \in \mathscr{W})$ is a trimming function useful for avoiding extreme values and $\xi_\tau(\cdot)$ is the check function defined earlier. As discussed in Lee (2007), the minimisation problem in (9) can be easily solved by computation methods developed for quantile regression models since it has a linear programming representation. The two-step quantile estimator $(\hat{\alpha}_x(\tau),\ \hat{\alpha}'_z(\tau))$ is asymptotically normal and the variance-covariance matrix can be consistently estimated through kernel methods that lead to valid asymptotic inferences.

In our case, the dependent variable i_t, is the interest rate. We set $x_t \equiv epu_t$ and $z_{1t} \equiv (\pi_{t+12}, y_t)'$, where epu_t is the economic policy uncertainty, π_{t+12} is the forward looking inflation[9] and y_t denotes the output gap.

3 Data and empirical analysis

3.1 Data

Our analysis uses monthly data on four countries (regions) namely the Euro Area (EA), Japan, the United Kingdom (UK) and the United States (US), with the period of coverage being: 1995:01–2017:05.[10] While the start date is driven by the common starting- point of data across the four countries (regions) for the measure of interest rate, the end-point is purely driven by availability of data at the time the paper was being written. The estimation of the Taylor-rule involves four variables, namely a measure of output, inflation, interest rate and uncertainty. In our case, we use industrial production (seasonally-adjusted) and Hodrick-Prescott filter to create the measure of output gap, while month-on-month inflation is computed based on the (seasonally-adjusted) Consumer Price Index (CPI). Data on both these variables for all the four economies under consideration is derived from the main economic indicators (MEIs) database of the Organisation for Economic Co-operation and Development (OECD).

Besides using a quantiles-based estimation approach, a unique feature of our study is that our data sample covers both the conventional and unconventional monetary policy periods, with the latter ensuing in the wake of the zero lower bound (ZLB)-situation of the monetary policy instrument in these economies, following the global financial crisis of 2008. The myriad unconventional monetary policies (such as large scale asset purchases, a maturity extension program and efforts of forward guidance in order to manage expectations of a prolonged period of low policy rates) were all directed towards improving financial conditions for firms and thereby eventually supporting an expedited recovery from the financial crisis. Given the ZLB, and the fact that a range of unconventional monetary policies were pursued, for estimation of policy rules across the conventional and unconventional regimes of central banking, we would need a uniform and coherent measure of the monetary policy stance. For our purpose, we use the shadow short rate (SSR). The SSR is the nominal interest rate that would prevail in the absence of its effective lower bound. The SSR used in this paper is developed by Krippner (2013), based on (two-factor) models of term-structure, at a daily frequency for the four economies of our concern, and is available for download from the website of the Reserve Bank of New Zealand.[11] The yield curve-based framework developed by Krippner (2013) essentially removes the effect that the option to invest in physical currency (at an interest rate of zero) has on yield curves, resulting in a hypothetical "shadow yield curve" that would exist if physical currency were not available. The process allows one to answer the question: "what policy rate would generate the observed yield curve if the policy rate could be taken negative?" The "shadow policy rate" generated in this manner, therefore, provides a measure of the monetary policy stance after the actual policy rate reaches zero. The main advantage of the SSR is that it is not constrained by the ZLB and thus allows us to combine the data from the ZLB period with the data from the non-ZLB era, and use it as the common metric of monetary policy stance across the conventional and unconventional monetary policy episodes. Note that, to match the monthly frequency of our other variables, the end of the month value of the daily SSR is used for our analysis.[12]

Our final variable that goes in the estimation of the monetary policy rules is uncertainty. But, uncertainty is a latent variable, and in order to quantify the impact of uncertainty on the interest rate (SSR) decisions, one requires ways to measure the former. In this regard, besides the various alternative measures of uncertainty associated with financial markets (such as the implied-volatility indices (popularly called the VIX), realized volatility, idiosyncratic volatility of equity returns, corporate spread associated), primarily three broad approaches to quantify uncertainty exists (Gupta, Lau and Wohar, 2018) forthcoming: (1) A news-based approach, with the main idea behind this method being to perform searches of major newspapers for terms related to economic and policy uncertainty (EPU) and to use the results of this search to construct measures of uncertainty; (2) Measures of uncertainty from estimates of various types of small and large-scale structural models related to macroeconomics and finance. Specifically speaking, the uncertainty measure is the average time-varying variance in the unpredictable component of a large set of real and financial time-series, i.e. it attempts to capture the average volatility in the shocks to the factors that summarize real and financial conditions, and; (3) Uncertainty derived from dispersion of professional forecaster disagreement.

As far as the metric of uncertainty is concerned, we use the first approach, i.e. news-based measure of Baker, Bloom, and Davis (2016), primarily due to the act that the measure does not require any complicated estimation of a large-scale model to generate it in the first place, and hence, is not model-specific. In this regard note that, for the US, to measure policy-related economic uncertainty, Baker, Bloom, and Davis (2016) construct an index from three types of underlying components. The first component quantifies newspaper coverage of policy-related economic uncertainty; the second component reflects the number of federal tax code provisions set to expire in future years, and; the third component uses disagreement among economic forecasters as a proxy for uncertainty. For the EU, Japan and the UK, the index is based on only the first component.[13] Besides not requiring pre-estimation and suffering from the generated regressor problem to yield biased estimates (Chen

2015), the advantage of using such a measure of uncertainty is that the data is available publicly for download for multiple countries, including the ones of our concern.[14]

3.2 Estimates at the conditional mean

To fix ideas, Table 1 reports GMM estimates of the Taylor rule equation (1). Inflation, output gap and uncertainty measure are instrumented using appropriately chosen lags of these variables. The set of instruments are determined by choosing lags that are sufficiently long to avoid correlation with the error term. We use the *J*-test (Hansen 1982) for the validity of overidentifying restrictions for each set of chosen instruments. We also use the C-statistic (Hayashi 2000) to test for endogeneity.

Table 1: Forward looking monetary rule: conditional mean GMM results.

	α_{π}	α_{y}	α_{epu}	H0: epu is exogenous	H0: inflation is exogenous	C-statistic H0: gap is exogenous	J-statistic
EA	0.4496	0.2076***	−0.0911***	44.5158 [0.0000]	5.5599 [0.0184]	0.4899 [0.4840]	5.7363 [0.0568]
Japan	−0.1881	0.0373*	−0.0019	0.0042 [0.9480]	0.4788 [0.9235]	0.0370 [0.8473]	0.2092 [0.6474]
UK	0.1646	0.3096***	−5.0639***	28.8238 [0.0000]	3.3941 [0.0654]	0.7502 [0.3864]	2.0996 [0.3500]
US	1.5440***	0.1065	−0.0854***	15.3349 [0.0001]	3.5947 [0.0580]	3.7199 [0.0538]	5.2465 [0.0726]

Forward looking monetary rule: $i_t = \alpha_0 + \alpha_{\pi}\pi_{t+12} + \alpha_y y_t + \alpha_{epu} epu_t$, where i_t is the end of month shadow interest rate, π_t is inflation, y_t is the output gap, and epu_t is the news-based economic policy uncertainty index of Baker, Bloom, and Davis (2016). Instruments: 1–3 lags of inflation, output gap and epu. J-statistic and C-statistic stand for Hansen's (1982) test for the validity of overidentifying restrictions and difference-in-Sargan exogeneity test (Hayashi, 2000), respectively. Figures in square brackets are p-values.

The specification for equation (3) allows for a forward-looking rate of inflation 12 months ahead, $k = 12$ for inflation,[15] a contemporary output gap (the dependence of these countries monetary policy on current rather than expected output gaps agrees with general consensus as for example in Chevapatrakul, Kim, and Mizen (2009)), and a contemporaneous uncertainty measure. We have tried various lags of uncertainty measure and we report the contemporaneous effect since it is of most importance. *, **, *** denote statistical significance at 10%, 5% and 1%, respectively. Our results show that the set of instruments includes a constant, 1-3 lagged values of inflation, output gap and uncertainty measure. The inflation effect is statistically significant satisfying the "Taylor principle" that inflation increases trigger an increase in the real interest rate of 0.5 percent when inflation increases by 1 percentage point and the response to uncertainty is significantly negative, confirming the Brainard attenuation principle. We found no significant reaction to the output gap for the US. This result echoes Chevapatrakul, Kim, and Mizen (2009) results for the US. The inflation coefficient estimates for the EA, Japan and the UK are insignificant. The response to output gap is significant for the EA and the UK and the response to uncertainty confirms a particularly high negative response for the UK while we detect no significant reaction for Japan.[16]

3.3 The Taylor Rule at various quantiles

Table 2 reports the estimated coefficients at each quantile using Lee (2007) IVQ based on the conditional quantile distribution of interest rate regression. We use an equation with the same set of instruments and the same forward-looking horizon as reported above in Section 3.2. We also estimate unconditional quantile regressions using the methodology of Firpo, Fortin, and Lemieux (2009) to further investigate the response of central bankers to inflation when the zero lower bound is approached. The results are reported in Table 2 (see figures in parentheses).

Table 2: Forward looking Taylor-type rule: IVQ methodology (Lee 2007) results using 2nd-order power series.

	Quantiles											Wald test
	0.05	0.10	0.20	0.30	0.40	0.50	0.60	0.70	0.80	0.90	0.95	
Panel A: Inflation coefficients α_π												
EA	1.1137**	0.1782**	0.2634*	0.2096**	0.1184**	0.3404**	0.2882**	0.2488***	0.2252**	0.3968***	0.0078**	1.87
	(1.2969)	(1.4104)	(1.6923)	(0.5788)	(0.1753)	(0.0911)	(0.1121)	(−0.0876)	(0.0405)	(−0.0014)	(0.0228)	[0.0481]
Japan	0.1838	−0.7823	−1.0112	−0.9113**	−0.7776**	−0.9328	−0.3662	0.2545	0.0866	0.1024	0.1083	0.79
	(0.2618)	(−0.8390)	(−0.7305)	(−0.4512)	(−0.2589)	(−0.4783)	(−0.1454)	(0.3851)	(−0.0087)	(0.4277)	(0.4851)	[0.6379]
UK	1.1083	0.4742**	0.1664**	0.2092***	0.2994**	0.2040	0.1246	0.2449	0.4153	0.5465	0.6688***	0.71
	(1.4736)	(0.4708)	(0.5134)	(0.1841)	(0.1412)	(0.1752)	(0.1609)	(−0.0301)	(0.1050)	(−0.1552)	(0.0344)	[0.7103]
US	1.1323	1.4597***	0.8343***	0.1478***	0.7185***	1.2776***	1.4844***	1.8906***	2.1985***	2.4423***	2.6881***	3.79
	(−0.2150)	(3.3460**)	(1.5373)	(1.9153*)	(1.8270)	(2.1825***)	(1.5858***)	(1.2287**)	(1.7679***)	(1.5909***)	(1.8273**)	[0.000]
Panel B: Gap coefficients α_y												
EA	0.2667***	0.2944***	0.2675***	0.2805***	0.2681***	0.2318***	0.1909***	0.1611***	0.1601***	0.1572***	0.1388***	1.64
	(−0.0708)	(0.0068)	(0.0923)	(0.2305)	(0.3037***)	(0.3287***)	(0.3374***)	(0.3536***)	(0.3515***)	(0.2220***)	(0.1289***)	[0.0972]
Japan	0.1666	0.2160	0.0471***	0.0578***	0.0552**	0.0281***	0.0508***	0.0324***	0.0254**	0.0200**	0.0120***	1.72
	(0.0175)	(0.0557)	(0.0224)	(0.0148)	(0.0231)	(0.0111)	(0.0419)	(0.0869***)	(0.0701***)	(0.0319*)	(0.0095)	[0.0763]
UK	0.3635***	0.3141**	0.3113**	0.3219***	0.2907***	0.2674***	0.2676***	0.1835***	0.1883***	0.2454***	0.2692***	1.19
	(1.3480**)	(0.3128)	(−0.2645)	(0.1554)	(0.5824**)	(0.4337***)	(0.3379***)	(0.3417***)	(0.1999**)	(−0.0067)	(−0.0495)	[0.2963]
US	−0.4042	−0.2878	0.0074***	0.0185***	0.0495***	0.0597***	0.2080***	0.3070***	0.3638***	0.5211***	0.5512***	6.84
	(−0.4737**)	(−0.5905**)	(−0.0757)	(0.5699***)	(0.3727*)	(0.3401**)	(0.2260**)	(0.2513***)	(0.4159***)	(0.3124***)	(0.2539***)	[0.0000]
Panel C: EPU coefficients α_{epu}												
EA	−0.0935***	−0.0878***	−0.0831***	−0.0791***	−0.0782***	−0.0760***	−0.0727***	−0.0686***	−0.0591***	−0.0570***	−0.0522***	8.35
	(−0.1042***)	(−0.0619)	(−0.0724***)	(−0.0678***)	(−0.0477***)	(−0.0360***)	(−0.0250***)	(−0.0228***)	(−0.0191***)	(−0.0147***)	(−0.0086***)	[0.0000]
Japan	−0.0153***	−0.0306***	−0.0341***	−0.0341***	−0.0310***	−0.0385***	−0.0145***	0.0011***	−0.0005***	0.0001***	−0.0001***	0.73
	(−0.0035)	(0.0015)	(−0.0020)	(−0.0083**)	(−0.0038)	(−0.0038)	(−0.0006)	(−0.0005)	(−0.0036)	(−0.0034)	(−0.0029*)	[0.6923]
UK	−7.3442***	−6.9726***	−6.5009***	−6.4039***	−6.3889***	−6.3196***	−6.1370***	−6.0838***	−5.9039***	−5.2520***	−5.0515***	4.96
	(−6.6236***)	(−6.6829***)	(−3.8079***)	(−5.2878***)	(−7.4995***)	(−4.3608***)	(−3.3763***)	(−2.6225***)	(−2.2312***)	(−1.2849***)	(−0.7850***)	[0.0000]
US	−0.1171***	−0.0946***	−0.0958***	−0.1128***	−0.1070***	−0.0955***	−0.0791***	−0.0564***	−0.0384***	−0.0271***	−0.0078***	20.08
	(−0.0903***)	(−0.1205***)	(−0.0933***)	(−0.0994***)	(−0.1082***)	(−0.0681***)	(−0.0343***)	(−0.0064)	(0.0109**)	(0.0075)	(−0.0029)	[0.0000]

Monetary rule: $i_t = \alpha_0(\tau) + \alpha_\pi(\tau)\pi_{t+12} + \alpha_y(\tau)y_t + \alpha_{epu}(\tau)epu_t$, where i_t is the end of month shadow interest rate, π_{t+12} is the forward looking inflation, y_t is the output gap, and epu_t is the news-based economic policy uncertainty index of Baker, Bloom, and Davis (2016). Instruments: 1–3 lags of inflation, output gap and epu. *, **, *** indicate statistical significance at 10%, 5% and 1% statistical levels, respectively. Wald statistic tests the null hypothesis of parameter equality across conditional quantiles. IVQ regressions calculations use power series with $k = 2$. Figures in parentheses are unconditional quantile regression estimates (Firpo, Fortin and Lemieux, 2009). Figures in square brackets are p-values.

3.3.1 The case of the US

The coefficient on the inflation rate is significantly different from zero from the 10th to the 95th quantile. The response is greater than unity above the 50th quantile, supporting the Taylor principle, meaning that the Fed raises the policy rate by more than 1 percent point when faced with a 1 percent point increase in inflation. The responsiveness of the interest rate toward inflation at the upper end of the conditional distribution of interest rate suggests that the Fed responded more aggressively to higher levels of inflation and hence shows evidence of a deflation bias to monetary policy. In fact, the response to inflation is greater than 1.50 above the 70th conditional quantile distribution of the interest rate and generally increases across quantiles. In contrast, we do not observe evidence of aggressive response as the zero and lower boundsare approached in the US based on the conditional regression method. We investigate more precisely the Fed's reaction at the lower quantiles of interest rates based on the unconditional regression method following Firpo, Fortin, and Lemieux (2009). This method has the advantage that it analyses the specific reaction of the central banker when interest rate is at zero and lower bounds specifically. Table 2 shows that there is very aggressive significant reaction at the 10th unconditional quantile distribution of interest rate, with the Fed raising the Fed's fund rate by more than 3.35 percentage points when faced with a 1 percent point increase in inflation. A result that tends to support the proposition emphasized in the introduction that forward looking policy makers in an attempt to prevent deflation close to the zero lower bounds when faced with demand shocks, will react strongly to inflation so that expected inflation will fall less since the private sector will take the policy makers' actions into account. We further observe that the response to inflation remains insignificant (at the 5% level of significance) up to the 40th unconditional quantile distribution of the interest rate, with the response becoming more significant at higher levels of interest rate (and possibly inflation as well), hence confirming evidence of a deflationary bias.

The conditional regression method shows that the response to output gaps is significant from the 20th quantile onwards, with a stronger response to the output gap at the higher conditional distribution of interest rate (and possibly high inflation) and the magnitude is very close to the value of 0.5, advocated by Taylor (1993). The response to outputis insignificant at low conditional interest rates (and possibly low inflation). One could possibly argue that the Fed places high importance on inflationary pressures of output during periods of rising inflation.[17]

With respect to the impact of uncertainty on policy, the results indicate that the Fed has been concerned about volatility throughout much of the sample period with a however higher negative response at the conditional lower bounds when the uncertainty about the conditions in financial markets was important to the interest rate setting behavior from 2008 for a prolonged period of time. We also notice that such reaction dampens once the economy is away from the conditional lower interest rate bounds. These parameters range from −0.1171 to −0.0078. The unconditional regression method further confirms these results of negative response to uncertainty at the lower levels of interest rates, including its zero and lower bounds. So what might not have been an evidence of aggressive response to inflation as the conditional zero and lower bounds are approached in the US might well have been translated instead into the decreased reaction to uncertainty as financial markets predicament unfolded.

To this end, our results go to complement the findings of Chevapatrakul, Kim, and Mizen (2009) who found no detectable evidence of increasing aggression to inflation as the zero lower bound is approached in the US and Japan together with the following two additional results. Firstly, based on the unconditional regression method, we were able to discern that there has been aggressive reaction to inflation from the monetary authorities as the lower bound is approached and secondly, that the Fed decreases its reaction to volatility, which are largely consistent with the Brainard (1967) attenuation principle. This decreased reaction to uncertainty as we approach the lower bound translated into a higher cut in the interest rate to expand the economy.[18]

3.3.2 International evidence

To highlight the importance of the uncertainty measure and how it contributed to monetary policy instance, we conduct both the conditional and unconditional regression estimations for the EA, Japan and the UK, and report the results in Table 2 as well, along with that of the US. First, we find a general consensus based on the conditional quantile regression that uncertainty decreases the monetary authorities' reaction which are largely consistent with the Brainard (1967) attenuation principle, and the proposition of cautious policy under uncertainty by Blinder (1999), suggesting that the monetary policy maker becomes less aggressive to a particular variable when it becomes more uncertain. Second, we also notice an across the board greater cut in policy rates at lower conditional quantiles of interest rate, meaning that as one approaches the conditional lower bound, the central bankers have an incentive to cut the interest rate even more to ensure proper expansionary mechanism in the economy. While the magnitude of the EPU estimates of the EA are roughly in line with the US, we note a particularly high response of this parameter for the UK, most probably highlighting the fact of a small open economy undertaking quantitative easing in the face of increased risk to its highly exposed financial markets. The Bank of Japan reaction to uncertainty is also in line with higher negative response at lower conditional interest rate quantiles. The unconditional regression results support these results of higher response at the lower bound, except for Japan, whereby most of the estimates turn out to be insignificant. Hence the prevalence of nonlinearity in the reaction to uncertainty by virtually all the central banks lend support to the added information obtained from quantile estimates over estimates at the conditional mean.

Turning to the response to inflation for the EA, Japan and the UK, we notice that all the parameter estimates are insignificant based on both conditional and unconditional quantile regressions except for the EA, whereby the conditional regression estimates are significant but however not adhering to the Taylor principle with the exception at the conditional lower bound (5^{th} quantile). The response to the output gap is significant for all these three economies based on the conditional regression method (with parameter estimates well below the 0.5 magnitude advocated by Taylor (1993) for the US) with the unconditional regression estimates being mostly insignificant.

3.3.3 Robustness analysis

We assess the sensitivity of our findings. Firstly, the identification of the regression quantile coefficients is crucially based on the control function approach. Therefore we implement a robustness check on the functional form by complementing Table 2 conditional quantile regression results which is based on the second-order polynomial series, with third-order and fifth-order polynomials. The results are provided in Tables 3 and 4 in the Appendix.[19] We find that the qualitative robustness of the quantile estimates does not change that much with respect to the response of the Fed vis-a-vis the inflation rates for the third-order polynomial series and to some extent for the fifth-order polynomials. Whereas for the EA, Japan and the UK response to inflation qualitatively are the same for the third-order polynomials, the EU's response to inflation are quantitatively insignificant at the fifth-order polynomial. With regard to the response to output gaps, the results are qualitatively similar for the third-order polynomial series compared to the second-order polynomial and some divergence with the fifth-order polynomial, especially for Japan, whereby the responses turned out to be insignificant at all quantiles. The impact of uncertainty on monetary policy remains significant and negative across all estimated control functions' polynomials with the same higher response at the conditional lower bounds. However this is less so for Japan with the fifth-order polynomial series whereby significance at the 10% level is achieved only across the 40^{th} to the 60^{th} and the 95^{th} conditional interest rate quantile.

Secondly, we report a number of robustness checks, viz., backward-looking rule where the nominal interest rate responds to *past* inflation rates. This follows Carlstrom and Fuerst (2000) that backward-looking Taylor rules ensure that the monetary reaction function used by the central bank does not introduce real indeterminacy and avoids sunspot fluctuations. We also report results based on contemporaneous monetary rule and lastly we report results using annualized inflation rates. All the results are re-done using the conditional quantile regression based on the second-order control function polynomials. Table 5 in the Appendix reports the results on backward-looking monetary rule. The response to inflation for all economies are qualitatively the same based on forward looking monetary policy rule, with the response to inflation for Japan and the UK remaining mainly insignificant and the response by the Fed to inflation are slightly lesser in magnitude. The Taylor rule estimates based on contemporaneous inflation rate are presented in Table 6 in the Appendix and conforms to the results obtained using both the forward and backward looking models with however a response to inflation by the Fed that only adheres to the Taylor principle from the 70^{th} conditional interest rate quantile onwards. Table 7 in the Appendix provides the results with the annualized inflation rate based on monthly data. The response to inflation rate for the EA turns out to be insignificant for couple of interest rate quantiles and at times significant and with the wrong sign for Japan and the UK. It shows that the Fed's response to inflation is significant and greater than unity across all conditional interest rate quantiles. While the response to uncertainty remains negative and significant for the EA, the UK and the US, the results turn out to be insignificant at the 5% level of significance for Japan at all quantiles altogether.

4 Conclusion

This paper provides some new results to the conduct of conventional and unconventional monetary policy for advanced economies. Importantly, the paper bridges the gap between existing empirical studies such as Chevapatrakul, Kim, and Mizen (2009) that found no detectable evidence of increasing aggression as the zero lower bound is approached in economies such as the US and Japan and the strong theoretical prediction that central banks should guard against the possibility of deflation if they operate near lower bounds by being willing to act especially forcefully, which were done in practice with Quantitative Easing for instance. To tackle this issue, the empirical analysis conducted in this paper makes use of the shadow interest rates, i.e. the nominal interest rate that would prevail in the absence of its effective lower bound, together with complementing the conditional interest rate quantile regressions with unconditional quantile regressions which allow more precise statements as lower bounds of interest rate are approached. We further make use of a news-based measure of uncertainty in the aftermath of the recent financial crisis. The results show an increased negative reaction of the Fed and other central banks towards uncertainty particularly at lower conditional quantiles of interest rates, hence lending support to increased aggressiveness as the conditional lower bound is approached together with increased aggressiveness by the Fed to inflation as the unconditional lower bound of interest rate is approached.

Other results emerge for the US such as adherence to the Taylor principle not just at the mean of interest rate and inflation but also across the different quantiles of interest rate. The prevalence of nonlinearity is present with the Fed reacting more aggressively at higher conditional levels of interest rate when inflation increases. These results clearly show that relying on linear models to investigate monetary policy reaction functions might be misleading and furthermore measures of uncertainty is of particular importance to gauge monetary policy reaction especially in dire situations such as crisis times.

Funding

This work was supported by the Open University of Cyprus, Funder Id: http://dx.doi.org/10.13039/100012996, Grand No: 2018/20, Essays in Economics.

Appendix

Table 3: Forward Looking Taylor-type rule: IVQ (Lee 2007) results using 3rd-order power series.

										τ	**Wald test**
	0.05	0.10	0.20	0.30	0.40	0.50	0.60	0.70	0.80	0.90	0.95	
Panel A: Inflation coefficients α_π												
EA	0.9855**	0.5613**	0.1964*	0.1382**	−0.0531**	−0.0129**	−0.2096**	0.4300***	0.3431**	0.2855***	0.1680**	1.84 [0.0544]
Japan	−0.1245	−0.7521	−0.9478	−0.5560**	−0.4393**	−0.6678	−0.3321	0.2027	0.1027	0.1264	0.1334	0.64 [0.7821]
UK	0.5238	1.0574**	0.4919**	0.0701***	−0.0509**	−0.0432	−0.0562	0.2365	0.1296	0.4274	0.1338***	0.54 [0.8577]
US	1.1323	1.4597***	0.8343***	0.1478***	0.7185***	1.2776***	1.4844***	1.8906***	2.1985***	2.4423***	2.6881***	3.03 [0.0011]
Panel B: Gap coefficients α_y												
EA	0.0146***	0.0506***	0.2194***	0.2432***	0.2720***	0.2708***	0.2229***	0.2181***	0.2034***	0.2286***	0.2094***	3.49 0.0003
Japan	0.1856	0.2015	0.0674***	0.0514***	0.0702**	0.0227***	0.0591***	0.0257***	0.0243**	0.0213**	0.0240***	1.73 [0.0767]
UK	0.2063***	0.3578***	0.2922**	0.3227***	0.2056***	0.2775***	0.2697***	0.3099***	0.3731***	0.2441***	0.1445***	1.46 [0.1537]
US	−0.4042	−0.2878	0.0074***	0.0185***	0.0495***	0.0597***	0.2080***	0.3070***	0.3638***	0.5211***	0.5512***	4.33 [0.0000]
Panel C: EPU coefficients α_{epu}												
EA	−0.0828***	−0.0793***	−0.0590***	−0.0463***	−0.0429***	−0.0417***	−0.0401***	−0.0353***	−0.0249***	−0.0223***	−0.0223***	8.11 [0.0000]
Japan	−0.0369***	−0.0497***	−0.0301***	−0.0170***	−0.0101***	−0.0139***	−0.0042***	0.0025***	−0.0003***	0.0011***	−0.0020***	0.77 [0.6547]
UK	−5.9176***	−5.6592***	−5.1127***	−5.1449***	−5.0611***	−4.6806***	−4.8604***	−4.9421***	−4.1461***	−4.1799***	−3.8998***	2.61 [0.0050]
US	−0.1171***	−0.0946***	−0.0958***	−0.1128***	−0.1070***	−0.0955***	−0.0791***	−0.0564***	−0.0384***	−0.0271***	−0.0078***	14.61 [0.0000]

Monetary rule: where i_t is the end of month shadow interest rate, π_{t+12} is the forward looking inflation, y_t is the output gap, and epu_t is the economic policy uncertainty index. Instruments: 1–3 lags of inflation, output gap and epu. *, **, *** indicate statistical significance at 10%, 5% and 1% statistical levels, respectively. Wald statistic tests the null hypothesis of parameter equality across quantiles. IVQ regressions calculations use power series with $k = 3$. Figures in square brackets are p-values.

Table 4: Forward looking Taylor-type rule: IVQ (Lee 2007) results using 5th-order power series.

	τ											Wald test
	0.05	0.10	0.20	0.30	0.40	0.50	0.60	0.70	0.80	0.90	0.95	
Panel A: Inflation coefficients α_π												
EA	0.7156	0.6712	0.0660	0.1400	−0.1049	−0.1784	−0.1406	0.3669	0.2570	0.2908	0.0440	2.28 [0.0182]
Japan	0.3570	−0.4202	−1.0473*	−0.5144	−0.4767	−0.5719	−0.3461	0.1648	0.1684	0.1025	0.0439	0.64 [0.7821]
UK	0.6694	0.7858	0.5123	0.6291*	0.1803	0.1125	0.2017	−0.0681	0.1606	0.3624	−0.5817	0.58 [0.8317]
US	1.5754**	1.0651*	1.2215*	0.6256	0.6331	1.0380	1.2652*	1.8998***	2.0167***	2.4165***	1.5310**	1.94 [0.0389]
Panel B: Gap coefficients α_y												
EA	0.0465	0.0116	0.1801*	0.2346***	0.2698***	0.2440***	0.2334***	0.1945***	0.1899***	0.2056***	0.2293***	2.35 [0.0117]
Japan	0.1965	0.1963	0.0958	0.0737	0.0664	0.0308	0.0614*	0.0404*	0.0273*	0.0189*	0.0171*	1.40 [0.1796]
UK	0.0607	0.1224	0.2271**	0.2352***	0.2428***	0.2236***	0.2278***	0.2942***	0.3955***	0.2071	0.1453	1.82 [0.0588]
US	−0.5042***	−0.3328***	−0.1187	−0.0409	−0.0102	0.0831	0.2796**	0.3844***	0.5259***	0.5676***	0.7133***	5.45 [0.0000]
Panel C: EPU coefficients α_{epu}												
EA	−0.0797***	−0.0751***	−0.0592***	−0.0456***	−0.0430***	−0.0408***	−0.0413***	−0.0351***	−0.0261***	−0.0220***	−0.0193***	6.65 [0.0000]
Japan	0.0039	0.0156	−0.0063	−0.0084*	−0.0082*	−0.0075*	−0.0051*	0.0006	0.0009	0.0002	−0.0013*	0.77 [0.6547]
UK	−5.8716***	−5.7361***	−5.2243***	−5.1449***	−5.0436***	−5.0415***	−4.9331***	−4.8439***	−4.8017***	−4.4973***	−3.8969***	2.90 [0.0020]
US	−0.0802***	−0.0904***	−0.0912***	−0.0901***	−0.0904***	−0.0762***	−0.0564***	−0.0391***	−0.0087	0.0017	0.0036	26.03 [0.0000]

Monetary rule: $i_t = \alpha_0(\tau) + \alpha_\pi(\tau)\pi_{t+12} + \alpha_y(\tau)y_t + \alpha_{epu}(\tau)epu_t$, where i_t is the end of month shadow interest rate, π_{t+12} is the forward looking inflation, y_t is the output gap, and epu_t is the economic policy uncertainty index. Instruments: 1–3 lags of inflation, output gap and epu. *, **, *** indicate statistical significance at 10%, 5% and 1% statistical levels, respectively. Wald statistic tests the null hypothesis of parameter equality across quantiles. IVQ regressions calculations use power series with $k = 5$. Figures in square brackets are p-values.

Table 5: Backward-looking Taylor-type rule: IVQ (Lee 2007) results using 2nd-order polynomial.

	τ											
	0.05	0.10	0.20	0.30	0.40	0.50	0.60	0.70	0.80	0.90	0.95	
Panel A: inflation coefficients α_π												
EA	0.6978**	0.4473**	0.3162*	0.1867**	0.3488**	0.3469**	0.5015**	0.4372***	0.6058**	0.3232***	0.4689**	2.02 [0.0317]
Japan	−0.7498	−1.3574	−0.9381	−0.4288**	−0.0373**	0.2270	0.2098	0.5097	0.1833	0.0542	0.0442	0.99 [0.4534]
UK	−0.1041	0.2358**	−0.3443**	−0.4066***	−0.4695**	−0.6316	−0.4131	−0.2923	0.0544	−0.1206	−0.0682***	0.22 [0.9938]
US	0.3953	0.5544***	0.6996***	0.8303***	0.6530***	1.1338***	1.2956***	1.4050***	1.5551***	2.5220***	2.1431***	0.89 [0.5427]
Panel B: Gap coefficients α_y												
EA	0.3040***	0.3088***	0.2711***	0.2684***	0.2452***	0.2370***	0.2025***	0.1896***	0.1512***	0.0933***	0.0541***	0.80 [0.6316]
Japan	−0.2670	−0.0183	0.0662***	0.0302***	0.0212**	0.0367***	0.0434***	0.0587***	0.0293**	0.0233**	0.0257***	1.98 [0.0356]
UK	0.2060***	0.4369***	0.3385**	0.2961***	0.2908***	0.2934***	0.3527***	0.3985***	0.3174***	0.2847***	0.1630***	0.79 [0.0000]
US	−0.2820	−0.2431	0.0496***	0.4184***	0.0322***	0.0321***	0.2533***	0.3740***	0.6344***	0.5869***	0.5550***	2.91 [0.0016]
Panel C: EPU coefficients α_{epu}												
EA	−0.0898***	−0.0899***	−0.0841***	−0.0825***	−0.0810***	−0.0792***	−0.0777***	−0.0756***	−0.0750***	−0.0691***	−0.0730***	15.80 [0.0000]
Japan	−0.0780***	−0.0731***	−0.0643***	−0.0712***	−0.0603***	−0.0438***	−0.0331***	0.0073***	−0.0015***	0.0025***	−0.0037***	0.97 [0.4696]
UK	−6.0397***	−5.3121***	−4.9415***	−4.7365***	−4.4479***	−4.1389***	−3.8003***	−3.3262***	−3.2612***	−3.3176***	−3.8631***	2.60 [0.0053]
US	−0.0803***	−0.0903***	−0.0902***	−0.0938***	−0.0890***	−0.0810***	−0.0678***	−0.0491***	−0.0282***	−0.0033***	−0.0004***	8.97 [0.0000]

Monetary rule: $i_t = \alpha_0(\tau) + \alpha_\pi(\tau)\pi_{t-12} + \alpha_y(\tau)y_t + \alpha_{epu}(\tau)epu_t$, where i_t is the end of month shadow interest rate, π_{t-12} is the backward looking inflation, y_t is the output gap, and epu_t is the economic policy uncertainty index. Instruments: 1–3 lags of inflation, output gap and epu. *, **, *** indicate statistical significance at 10%, 5% and 1% statistical levels, respectively. Wald statistic tests the null hypothesis of parameter equality across quantiles. IVQ regressions calculations use power series with $k = 2$. Figures in square brackets are p-values.

Table 6: Contemporaneous Taylor-type rule: IVQ (Lee 2007) results using 2nd-order power series.

	τ											Wald test
	0.05	0.10	0.20	0.30	0.40	0.50	0.60	0.70	0.80	0.90	0.95	
Panel A: Inflation coefficients α_π												
EA	0.0983**	0.6857**	0.6691*	0.4013**	0.1192**	0.0211**	0.0334**	0.3481***	0.2995**	0.2288***	0.0231**	0.68 [0.7442]
Japan	−0.0976	−0.5565	−0.7669	−0.7158**	−0.5436**	−0.2532	0.1621	0.2267	0.0376	0.0231	0.0754	0.91 [0.5233]
UK	0.03831	0.1723**	0.0583**	−0.2631***	−0.3839**	−0.1409	0.2816	0.6047	0.1386	−0.0481	0.1586***	0.33 [0.9726]
US	1.6387	0.7714***	1.1919***	0.8855***	0.6120***	0.7107***	0.9847***	1.2719***	2.3995***	1.8278***	2.3486***	0.78 [0.6480]
Panel B: Gap coefficients α_y												
EA	0.1831***	0.0797***	0.1898***	0.2442***	0.2368***	0.2952***	0.2356***	0.2144***	0.1834***	0.2292***	0.2079***	1.64 [0.0950]
Japan	0.2032	0.0893	0.1860***	0.0878***	0.0663**	0.0121***	0.0681***	0.0321***	0.0246**	0.0179**	0.0079***	2.08 [0.0266]
UK	0.2649***	0.3495***	0.3248**	0.3175***	0.2795***	0.2485***	0.3199***	0.3908***	0.3516***	0.2984***	0.1347***	0.74 [0.6890]
US	−0.3247	−0.2405	0.0368***	0.0101***	0.0220***	0.0677***	0.3109***	0.4347***	0.3981***	0.5660***	0.6702***	3.87 [0.0001]
Panel C: EPU coefficients α_{epu}												
EA	−0.0815***	−0.0778***	−0.0731***	−0.0582***	−0.0458***	−0.0448***	−0.0431***	−0.0367***	−0.0282***	−0.0215***	−0.0210***	14.04 [0.0000]
Japan	−0.0101***	−0.0135***	−0.0226***	−0.0096***	−0.0093***	−0.0103***	−0.0047***	0.0017***	−0.0001***	0.0003***	−0.0015***	0.72 [0.7026]
UK	−5.7669***	−5.2337***	−4.9601***	−4.8114***	−4.5953***	−4.3952***	−4.0064***	−3.5966***	−3.6440***	−3.7882***	−3.8795***	4.25 [0.0000]
US	−0.0791***	−0.0814***	−0.0839***	−0.0882***	−0.0894***	−0.0794***	−0.0506***	−0.0368***	−0.0069***	0.0089***	0.0180***	13.09 [0.0000]

Monetary rule: $i_t = \alpha_0(\tau) + \alpha_\pi(\tau)\pi_t + \alpha_y(\tau)y_t + \alpha_{epu}(\tau)epu_t$, where i_t is the end of month shadow interest rate, π_t is the inflation, y_t is the output gap, and epu_t is the economic policy uncertainty index. Instruments: 1–3 lags of inflation, output gap and epu. *, **, *** indicate statistical significance at 10%, 5% and 1% statistical levels, respectively. Wald statistic tests the null hypothesis of parameter equality across qantiles. IVQ regressions calculations use power series with $k = 2$. Figures in square brackets are p-values.

Table 7: Forward looking Taylor-type rule with annualised inflation: IVQ (Lee 2007) results using 2nd-order power series.

	τ											Wald test
	0.05	0.10	0.20	0.30	0.40	0.50	0.60	0.70	0.80	0.90	0.95	
Panel A: Inflation coefficients α_{π}												
EA	0.7739***	1.2431***	1.2799***	0.7421***	0.5484**	0.3304*	0.0863	0.0739	0.1374	0.2447*	0.2620**	1.94 [0.0410]
Japan	−0.7488	−0.3195	−0.5923***	−0.6272***	−0.4285***	−0.4297***	−0.3278**	0.0928	0.0499	0.0483	0.0553	0.74 [0.6885]
UK	−0.1516	−0.0541	−0.1021	−0.0586	−0.0177	−0.0926	−0.1121	0.1828	−0.4259**	−0.9451***	−1.0793***	0.59 [0.4497]
US	1.0465***	0.9584***	1.1359***	1.3528***	1.5158***	1.5006***	1.1580***	1.0909***	1.1112***	1.1219***	1.0888***	1.79 [0.0609]
Panel B: Gap coefficients α_{y}												
EA	−0.0622	−0.0328	0.0714	0.1642*	0.2028**	0.2745***	0.2753***	0.2289***	0.1914***	0.2090***	0.2149***	1.83 [0.0561]
Japan	0.2392	0.2033	0.1227**	0.1002*	0.0692	0.0058	0.0331	0.0555**	0.0254*	0.0182*	0.0200**	1.69 [0.0822]
UK	0.0404	0.2817*	0.1834*	0.2621***	0.2103***	0.2030***	0.2425***	0.3086***	0.4397***	0.4334***	0.4385***	1.15 [0.3239]
US	0.0158	−0.1093	−0.2190*	−0.2168*	−0.1083	0.0671	0.2200**	0.3137***	0.3598***	0.3888***	0.1675*	9.39 [0.0000]
Panel C: EPU coefficients α_{epu}												
EA	−0.0625***	−0.0434***	−0.0343***	−0.0388***	−0.0379***	−0.0432***	−0.0423***	−0.0363***	−0.0279***	−0.0215***	−0.0199***	9.08 [0.0000]
Japan	0.0060	0.0056	−0.0105*	−0.0092*	−0.0088*	−0.0094*	−0.0109*	0.0035	0.0006	−0.0001	−0.0000	0.80 [0.6256]
UK	−5.7971***	−5.4694***	−5.0885***	−4.9765***	−4.9029***	−4.6028***	−4.5861***	−4.4474***	−3.3961***	−3.3852***	−3.4977***	3.21 [0.0007]
US	−0.0671***	−0.0734***	−0.0769***	−0.0780***	−0.0784***	−0.0706***	−0.0586***	−0.0473***	−0.0400***	−0.0289***	−0.0016***	20.66 [0.0000]

Monetary rule: $i_t = \alpha_0(\tau) + \alpha_{\pi}(\tau)\pi_{t+12} + \alpha_y(\tau)y_t + \alpha_{epu}(\tau)epu_t$, where i_t is the end of month shadow interest rate, π_{t+12} is the forward looking annualised inflation, y_t is the output gap, and epu_t is the economic policy uncertainty index. Instruments: 1–3 lags of inflation, output gap and epu. *, **, *** indicate statistical significance at 10%, 5% and 1% statistical levels, respectively. Wald statistic tests the null hypothesis of parameter equality across quantiles. IVQ regressions calculations use power series with $k = 5$. Figures in square brackets are p-values.

Notes

1 Cukierman and Gerlach (2003), Ruge-Murcia (2003), Dolado, María Dolores, and Ruge-Murcia (2004), and Dolado, María-Dolores, and Naveira (2005)), and Surico (2007) have shown evidence supporting asymmetries in nonlinear monetary policy reaction function for major OECD economies and Castro and Sousa (2012) assess the response of monetary policy to developments in asset markets in the euro area, the UK and the US. While Jawadi, Mallick, and Sousa (2014) provide support for asymmetric monetary policy reaction for emerging market economies such as Brazil and China.

2 Quantiles-based estimation of Taylor rules for emerging and developing Asian and Latin American countries can be found in Miles and Schreyer (2012, 2014), and Christou et al. (forthcoming).

3 This literature presents evidence in support of the attenuation principle, viz., Svensson (1999), Peersman and Smets (1999), Estrella and Mishkin (2000), Orphanides et al. (2000), Rudebusch (2001), Ehrmann and Smets (2003), Martin and Costas Milas (2009), and Naraidoo and Raputsoane (2015). On the contrary, Giannoni (2002) and Sonderstrom (2002), among others, have presented evidence that supports an aggressive reaction of monetary policy under uncertainty. More recently, Caggiano, Castelnuovo, and Nodari (2018) and Ma, Olson, and Wohar (2018) found strong evidence indicating that the Fed responded to increases in financial and macroeconomic uncertainty respectively, by cutting the Federal Funds rate. The fact that the Fed did respond to macroeconomic uncertainty in the pre-zero lower bound (ZLB) period has also been earlier confirmed by Evans et al. (2015), when estimating modified Taylor-type monetary policy rules. Note that the negative response of interest rate to uncertainty (shocks) is also consistently detected in the vector autoregressive model-based analysis of uncertainty, where by uncertainty shocks are identified as negative demand shocks (see for example, Bloom, 2009; Colombo, 2013; Jones and Olson, 2015; Jurado, Ludvigson and Ng, 2015; Gupta, Lau and Wohar, 2018; and references cited there in).

4 Note that quantile regressions have been widely used in wide variety of applications in economics and empirical finance. In this regard, see for example, Taylor (1999), who considers quantile regressions in the context of value at risk; Conley and Galenson (1998), who explore wealth accumulation in several US cities, while Gosling, Machin, and Meghir (2000) study income and wealth distribution in the UK; Machado and Sousa (2005) assessed the impact of macroeconomic fundamentals on the distribution of asset prices; Leon Li and Yen (2011) analysed the dynamic covariance risk in global stock markets; Jawadi and Sousa (2013, 2014), respectively, estimated money demand equations (for the euro area, the US and the UK), and modelled the relationship between consumption and wealth (for the same economies); and Sousa and Sousa (2017) assessed time-variation in US asset returns while considering the whole conditional distribution, and show that the probabilistic distribution of expectations about future stock returns changes in response to variation in commonly used explanatory variables.

5 In this regard, an important study analysing the impact of unconventional monetary policy on the US economy is that by Jawadi, Sousa, and Traverso (2017). These authors investigated the macroeconomic and wealth effects of nonstandard monetary policy shocks using a Bayesian structural vector autoregression (B-SVAR) and US monthly data for the post- Lehman Brothers' collapse period. They show that such shocks do not have a substantial impact on industrial production or consumer prices, but are responsible for a strong boost to asset prices, which is larger in magnitude for stock prices than for housing prices.

6 In this regard note that, more recently, Agnello et al. (2018) specify unconventional monetary policy reaction functions for the Fed using a Markov-Switching Regression (MSR) and a Time-Varying Probability Markovian Process (TVPMS), which in turn can also be considered as an alternative approach to quantile regressions in modelling nonlinear Taylor-type rules.

7 A number of authors estimate a monetary rule with interest rate smoothing. However, in our case, all interest rate time series appear to be near to unit root processes. It is well known that values of interest rate smoothing parameter in the vicinity of unity cause parameter estimates to diverge (Chevapatrakul, Kim and Mizen, 2009). To solve this problem, we follow Chevapatrakul, Kim, and Mizen (2009) and we adopt a Taylor-type rule without interest rate smoothing.

8 In this paper we use power functions. Specifically, we provide results for second-third- and fifth-order polynomials.

9 Following the literature, we use the observed 12 month-ahead inflation.

10 As a word of caution, we must point out that, due to data availability, the sample period is 1995:01–2017:05, which is relatively short. Thus, estimating quantile regressions using such time window implies that the results are based on a very few data points even if all the distribution is used in the econometric framework. However, tests of equality of slope parameters across the quantiles are found to be different in a statistically significant manner, and hence, is an indication that the quantiles in our case have been estimated appropriately, and also captures the inherent nonlinearity in the relationship.

11 The data can be downloaded from the following link: https://www.rbnz.govt.nz/research-and-publications/research-programme/additional-research/measures-of-the-stance-of-united-states-monetary-policy/comparison-of-international-monetary-policy-measures.

12 For comparability, all of the estimates are obtained using the Krippner (2013) shadow/lower bound framework with two factors, i.e. the K-ANSM(2), a fixed 12.5 basis point lower bound, and yield curve data with maturities from 0.25 to 30 years with the sample beginning in 1995. SSR estimates can be very sensitive to the model specification, data, and estimation method. Krippner (2013) approach is designed to be as comparable as possible by holding each of the above aspects consistent between the four economies. In addition, SSR results from different K-ANSM(2) applications are shown by Krippner (2013) to be robust in profile and magnitude, and correspond well with unconventional monetary policy events. These properties do not generally hold for SSR estimates from three-factor models, which includes K-ANSM(3) model of Wu and Xia (2016). Hence our decision to use this database, besides the fact that Wu and Xia (2016) does not report SSR estimates for Japan.

13 However, when we use the first (i.e. news-based) component only of uncertainty for the US in our estimation, our results were qualitatively similar to those reported for the broad measure. Complete details of these results are available upon request from the authors.

14 The data can be downloaded by following the appropriate country (region)-specific links at: http://www.policyuncertainty.com/.

15 We also report results based on contemporaneous and backward looking monetary rules.

16 The estimates based on contemporaneous and backward looking monetary rules yield poor insignificant reaction to inflation while the reaction to output gap and uncertainty are qualitatively the same compared to the forward looking rule.

17 The unconditional regression shows significant reaction to output gaps at all quantiles with the wrong sign however at the 5th and 10th unconditional quantiles of interest rate.

18 We conducted additional robustness analyses to assess the sensitivity of our findings. First, we used a different inflation horizon (i.e. one-month ahead); second, we utilized the average over the month (instead of the end-of-month) measure of shadow short rate; third, we repeated the analysis using the narrower news-based measure of uncertainty, and; finally; we also investigated the quantile interest rate rule estimation using the new method suggested by Lee (2016) which also takes into account the persistence of the regressors. In general, our results were qualitatively similar to those reported in the paper, and are available upon request from the authors.

19 We would like to thank Lee (2007) for providing the codes used in our programming.

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Supplementary Material: The online version of this article offers supplementary material (DOI: https://doi.org/10.1515/snde-2018-0056).