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Asymmetric Effects of Policy Uncertainty on the Demand for Money in the United States [†]

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Abstract: A comprehensive measure of economic uncertainty, known as "Policy Uncertainty", which was constructed by the Economic Policy Uncertainty Group by searching popular newspapers for uncertain terms associated with economic factors and its impact on macro variables, is gaining momentum. Although some researchers have assessed its impact on the demand for money in a few countries, we considered the U.S.A. demand for money one more time and showed that when a linear money demand was estimated, policy uncertainty had no long-run effects. However, when a nonlinear model was estimated, the results showed that while increased policy uncertainty induces the public to hold less money in the long run, decreased uncertainty has no long-run effects, a clear sign of asymmetric response.

1. Introduction

Ever since the introduction of the new measure of uncertainty by Baker et al. (2016), response of macro variables to this new measure has gained momentum. Unlike other measures that are based on a single macro variable, such as volatility of money supply or output, under the new measure, which is known as "Policy Uncertainty", the Policy Uncertainty Group today relies upon the method by Baker et al. (2016) and constructs the policy uncertainty measure by searching popular newspapers in a given country for such terms as "uncertain' and "uncertainty" associated with such words as "policy", "tax", "spending", "regulation", "central bank", "budget", "deficit", etc. From the volume of news associated with these terms, an index of uncertainty is then constructed. The larger the volume of the news, the higher the index, and the higher uncertainty.[1]

As mentioned, the new measure has recently gained momentum and researchers are emphasizing its impact on macro variables. Examples include Wang et al. (2014) who investigated the link between policy uncertainty and corporate investment, Pastor and Veronesi (2013); Ko and Lee (2015); and Brogaard and Detzel (2015), who assessed its impact on risk premia and market returns, Baker et al. (2016) who assessed its impact on economic activity and firm-level outcome, Bahmani-Oskooee and Ghodsi (2017), who investigated its impact on house prices in each state of

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[1] For more information and source of the data visit Economic Uncertainty Policy Group: http://www.policyuncertainty.com/europe_monthly.html.

the U.S.A., and Kang and Ratti (2013), as well as Bahmani-Oskooee et al. (2018), who looked into the link between the new uncertainty measure and oil prices.

Another macro variable that is said to be affected by an uncertainty measure is the amount of money that people hold as cash, i.e., the demand for money. Assessing the impact of an uncertainty measure on the demand for money must be first attributed to Friedman (1984), who emphasized volatility of monetary growth as a measure of uncertainty. Choi and Oh (2003) then added volatility of output as a measure of uncertainty that could affect the demand for money. However, Bahmani-Oskooee et al. (2015, 2016) argued that factors other than money supply and output volatility can contribute to an uncertain economic environment. Hence, they employed the new policy uncertainty measure and assessed its impact on the demand for money in the U.K. and in the U.S.A., respectively.[2]

All of the above studies have assumed that the impact of policy uncertainty on macro variables is symmetric. However, concentrating on the demand for money, Bahmani-Oskooee and Maki-Nayeri (2018a, 2018b) recently argued that the effects of policy uncertainty on the demand for money could be asymmetric. As they argued, while people hold more cash during times of increased uncertainty, people might hold even more cash during times of decreased uncertainty, as they might attempt shield themselves from an uncertain environment in the future. They demonstrated these asymmetric effects by estimating the demand for money in Korea and Australia, respectively.

As mentioned above, Bahmani-Oskooee et al. (2016) have already assessed the impact of the new measure of policy uncertainty on the demand for money in the U.S.A., and have shown that the new uncertainty measure has short-run and long-run positive effects, implying that due to policy uncertainty Americans hold more cash so that they can hedge against an uncertain future. The findings also imply that a decrease in policy uncertainty will induce the U.S.A. public to hold less cash, implying a symmetric response to policy uncertainty. How valid is this symmetric assumption in the U.S.A.? As our asymmetry analysis will show, the assumption is valid in the short run, though not in the long run. Therefore, the main goal of this paper is to investigate the asymmetric effects of policy uncertainty on the demand for money in the U.S.A. To gain some insight on the path of policy uncertainty measure in the U.S.A., see Figure 1. The rest of the paper is organized as follows: in Section 2, we outline the models and explain the methods. We then present the estimation results in Section 3 and provide a summary in Section 4. Data definition and sources are then explained in Appendix A.
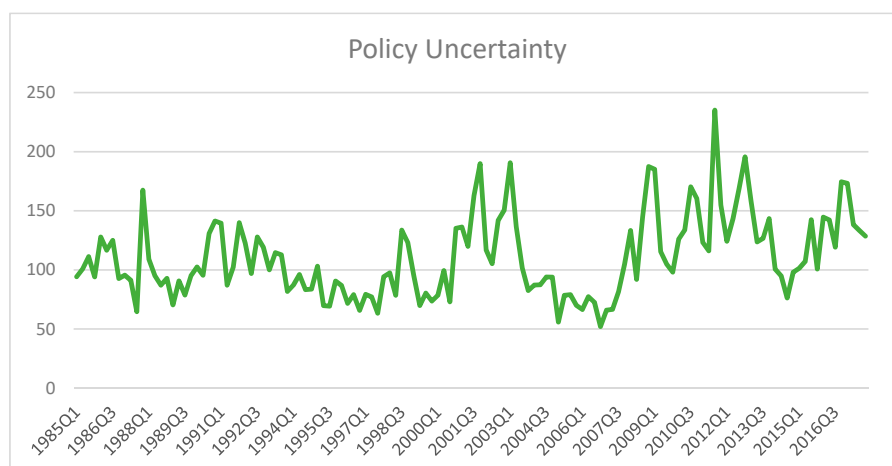
Figure 1. Measure of Policy Uncertainty for the U.S.A.

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[2] Other measures of uncertainty have been used by others. For example, Sahin (2013) assessed the impact of inflation uncertainty on the demand for money in Turkey and found that increased uncertainty increased precautionary motives. Sahin (2018) looked at the impact of impact of the U.S.A. money supply volatility as a measure of uncertainty on the velocity of the money in Organization of the Petroleum Exporting Countries (OPEC) countries and finds significant long-run effects only in three countries.

2. The Money Demand Models and Methods

Since our main purpose is to extend the symmetric analysis of Bahmani-Oskooee et al. (2015) and engage in an asymmetric analysis, we borrow their specification as outlined by Equation (1):

$$LnM_t = a + bLnY_t + cLnr_t + dLn(P_t/P_{t-1} + eLnEX_t + fLnPU_t + \varepsilon_t \tag{1}$$

In specification (1), it is basically assumed that level of real income, Y, interest rate, r, rate of inflation proxied by $Ln(P_t/P_{t-1})$, the nominal effective exchange rate, EX, and policy uncertainty, PU, are the main determinants of the demand for money in the U.S.A. Real income was included to account for transaction demand for money and it was expected that income elasticity to be positive. Interest rate was included to account for the opportunity cost of holding money against other financial assets and inflation rate was included as opportunity cost against real assets. Estimates of both c and d were expected to be negative. The nominal effective exchange rate was included to account for currency substitution, and as discussed by Bahmani-Oskooee et al. (2015) and others in the literature, an estimate of e could be either negative or positive. As the U.S. dollar depreciates (i.e., *EX* declines), domestic currency value of foreign assets rises; if this is perceived as an increase in wealth, foreign asset holders at home will increase their spending by holding more cash, hence a negative estimate for *e*. On the other hand, as the dollar depreciates or foreign currencies appreciate, some may expect further appreciation of foreign currencies. They will then hold more foreign currencies and fewer dollars, hence, a positive estimate for e.[3] Similarly, since policy uncertainty could have negative or positive impact on the demand for money, an estimate of f could be positive or negative.[4]

Coefficient estimates of model (1) are the long-run elasticities. In order to infer the short-run effects of exogenous variables, we rely upon an estimation method that yields both short-run and long-run effects in one step, i.e., the Autoregressive Distributed Lag (ARDL) approach of Pesaran et al. (2001). Accordingly, we turned (1) to an error-correction model as follows:

$$
\begin{aligned}
\Delta LnM_t = \quad & \alpha + \sum_{i=1}^{n1} \beta_i \Delta LnM_{t-i} + \sum_{i=0}^{n2} \delta_i \Delta LnY_{t-i} + \sum_{i=0}^{n3} \varphi_i \Delta Lnr_{t-i} + \sum_{i=0}^{n4} \gamma_i \Delta Ln(P_t/P_{t-1})_{t-i} + \\
& \sum_{i=0}^{n5} \eta_i \Delta LnEX_{t-i} + \sum_{i=0}^{n6} \lambda_i \Delta LnPU_{t-i} + \rho_0 LnM_{t-1} + \rho_1 LnY_{t-1} + \\
& \rho_2 Lnr_{t-1} + \rho_3 Ln(P_t/P_{t-1})_{t-1} + \rho_4 LnEX_{t-1} + \rho_5 LnPU_{t-1} + \varepsilon_t
\end{aligned}
\tag{2}
$$

As can be seen, once the error-correction model (2) is estimated, short-run effects of each variable are reflected in the estimate of coefficients attached to first-differenced variables. For example, short-run effects of policy uncertainty are inferred by the estimates of λ_i's. The long-run effects are inferred by the estimates of $\rho_1–\rho_5$ normalized on ρ_0. However, in order to avoid spurious regression problem, we applied the F-test to establish joint significance of lagged level variables as a sign of cointegration. Pesaran et al. (2001) tabulated new critical values for the F-test that accounted for integrating properties of variables. Indeed, variables could be a combination of I(0) and I(1), but not I(2) and this is another advantage of this approach.[5]

As mentioned before, our goal was to extend the above symmetric analysis to asymmetric analysis. To this end, we followed Bahmani-Oskooee and Maki-Nayeri (2018a) and Shin et al. (2014) and decomposed the *LnPU* variable into two new time-series variables. For this purpose, we first form

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[3]   For more, see Mundell (1963); Arango and Nadiri (1981); and Bahmani-Oskooee and Pourheydarian (1990).
[4]   Other studies that have estimated the demand for money in the U.S.A. without uncertainty measure are: Hafer and Jansen (1991); Hoffman and Rasche (1991); McNown and Wallace (1992); Ahking (2002); Wang (2011); Rao and Kumar (2011); Ball (2012); Jawadi and Sousa (2013); and Gupta and Majumdar (2014)Gupta and Majumdar
[5]   Another advantage of this approach is that by including short-run dynamic adjustment process in estimating the long-run elasticities, the approach accounts for feedback effects among all variables (Pesaran et al. 2001, p. 299).

$\Delta LnPU$ which includes positive changes as well as negative changes in our uncertainty measure. We then used the partial sum concept to construct the two new series as follows:

$$POS_t = \sum_{j=1}^{t} \max(\Delta LnPU_j, 0), \ NEG_t = \sum_{j=1}^{t} \min(\Delta LnPU_j, 0) \tag{3}$$

where POS_t, which is the partial sum of positive changes, reflects only increases in policy uncertainty and NEG_t, which is the partial sum of negative changes in uncertainty, reflects only declines in policy uncertainty. The next step is to move back to ARDL model (2) and replace $LnPU$ by POS and NEG variables. We then arrive at the following specification:

$$
\begin{aligned}
\Delta LnM_t = \ & \alpha + \sum_{i=1}^{n1} \beta_i \Delta LnM_{t-i} + \sum_{i=0}^{n2} \delta_i \Delta LnY_{t-i} + \sum_{i=0}^{n3} \varphi_i \Delta Lnr_{t-i} + \sum_{i=0}^{n4} \gamma_i \Delta Ln(P_t/P_{t-1})_{t-i} + \\
& \sum_{i=0}^{n5} \eta_i \Delta LnEX_{t-i} + \sum_{i=0}^{n6} \lambda_i^+ \Delta POS_{t-i} + \sum_{i=0}^{n7} \lambda_i^- \Delta NEG_{t-i} + \rho_0 LnM_{t-1} + \rho_1 LnY_{t-1} + \\
& \rho_2 Lnr_{t-1} + \rho_3 Ln(P_t/P_{t-1})_{t-1} + \rho_4 LnEX_{t-1} + \rho_5^+ POS_{t-1} + \rho_5^- NEG_{t-1} + \varepsilon_t
\end{aligned}
\tag{4}
$$

Since constructing the partial sum variables introduce nonlinearity into adjustment process, models like (4) are classified as nonlinear ARDL models (Shin et al. 2014), whereas (2) is labeled as the linear ARDL model. Shin et al. (2014) then demonstrated that the same estimation method and the same tests by Pesaran et al. (2001) could be applied to (4) as well. Indeed, they argue that the critical value of the F-test for cointegration should stay the same when we move from (2) to (4), even though (4) has one more variable.[6]

Once (4) was estimated, a few asymmetry assumptions could be tested. First, after using a specific lag selection criterion, if ΔPOS takes a different lag order than ΔNEG, that will be evidence of short-run adjustment asymmetry, implying that the public responds at a different speed to an increase in uncertainty versus a decrease in uncertainty. Second, at a given lag order such as $i = 1$, if estimate of $\lambda+$ differs from the estimate of λ^-, that will support short-run asymmetric effects of changes in policy uncertainty. However, stronger short-run cumulative or impact asymmetric effects will be established if we reject the null hypothesis of $\sum \lambda_i^+ = \sum \lambda_i^-$. The Wald test is commonly used to test this hypothesis. Finally, long-run asymmetric effects of policy uncertainty on the demand for money will be established if we reject the null hypothesis of $\frac{\rho_5^+}{-\rho_0} = \frac{\rho_5^-}{-\rho_0}$, again by applying the Wald test.[7]

3. The Results

In this section we estimate both the linear model (2) and the nonlinear model (4) using quarterly data from the U.S.A. over the period 1985I–2017IV. The main reason for restricting ourselves for this period was availability of data for policy uncertainty measure. In estimating each model we followed the literature and use Akaike's Information Criterion (AIC) to select the optimum number of lags. Furthermore, since there are different critical values for different estimates, they were collected in the notes to each table and were used to identify significant estimates. If an estimate was significant at the 10% level, one * is used. Those significant at the 5% level are identified by **.

One of the requirement of the ARDL bounds testing approach is to rule out potential I(2) variables. For this purpose, we applied the augmented Dickey-Fuller (ADF) test to the level, as well as first-differenced variables; the outcome is reported in Table 1 along with some descriptive statistics.

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6   See Shin et al. (2014, p. 219). This proposition is based on dependency between the two partial sum variables.
7   For some other application of these methods in recent literature, see Gogas and Pragidis (2015); Al-Shayeb and Hatemi-J (2016); Lima et al. (2016); Nusair (2017); Aftab et al. (2017); Arize et al. (2017); and Gregoriou (2017). Furthermore, we have used statistical package Microfit 5.5 by Pesaran and Pesaran downloadable for free at: http://www.econ.cam.ac.uk/people-files/emeritus/mhp1/Microfit/Microfit.html.

From Table 1, we gather that while inflation rate and policy uncertainty variable were both I(0), other variables were I(1). There were no I(2) variables.

Table 1. Summary Statistics for All Variables.

		Variables					
		M	*Y*	*r*	$Ln(P_t/P_{t-1})$	EX	PU
Mean		3273.31	12402.14	3.33	0.53	115.55	110.44
Min		2244.30	7537.93	0.01	−0.20	92.51	52.09
Max		5578.80	17272.47	8.54	1.20	170.40	235.08
Std Dev		976.73	2899.15	2.57	0.24	14.01	34.20
Skewness		0.90	−0.11	0.09	0.14	1.00	0.86
Kurtosis		2.62	1.65	1.72	3.52	4.74	3.59
ADF Test Results (Augmented Dickey-Fuller test)							
		Variables					
		Ln M	**Ln Y**	**Ln r**	$Ln(P_t/P_{t-1})$	**Ln EX**	**Ln PU**
With Constant	Level	−2.10(1)	−1.88(1)	−1.53(1)	−3.04(2) *	−2.88(2)	−5.22(0) **
	First Difference	−8.06(0) **	−4.88(1) **	−9.04(0) **	−12.33(1) **	−7.97(1) **	−12.22(1) **
With Constant and Trend	Level	−0.61(0)	−1.34(2)	−1.89(1)	−6.26(0) **	−2.99(1)	−5.54(0) **
	First Difference	−8.65(0) **	−7.78(0) **	−9.03(0) **	−12.28(1) **	−8.12(1) **	−12.17(1) **

Notes: Real money supply (M) is in Millions of U.S. Dollars. Std Dev is standard deviation. * and ** denote statistical significance at the 10% and 5% confidence levels, respectively. Number inside the parenthesis is the number of lags selected by AIC.

We are now in a position to estimate the linear ARDL model (2). The results are reported in Table 2. From the short-run results reported in Panel A, we gathered that all variables had short-run effects since each variable carried at least one significant coefficient. However, short-run effects of none of the variables last into the long run, since none of the normalized long-run estimates in Panel B were significant. This is supported by the lack of cointegration among the variables, since the F-test reported in Panel C was marginally significant at the 10% level and ECM_{t-1} was not.[8] The model seems to have been misspecified, since Ramsey's RESET test in Panel C was highly significant but there is no evidence of serial correlation, since the Lagrange Multiplier statistic reported as LM was insignificant. Finally, application of the CUSUM and CUSUMSQ tests for stability of all coefficient estimates reported in Figure 2 reveals that estimates were stable by the first test but not by the second test. These results are also indicated in Panel C by "S" for stable and "UNS" for unstable estimates. In sum, it appears that policy uncertainty has short-run effects on the demand for money in the U.S.A., though not long-run effects.

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[8]　The ECM_{t-1} test is an alternative test under which normalized long-run estimates and Equation (1) are used to generate the error term denoted by ECM as follows:
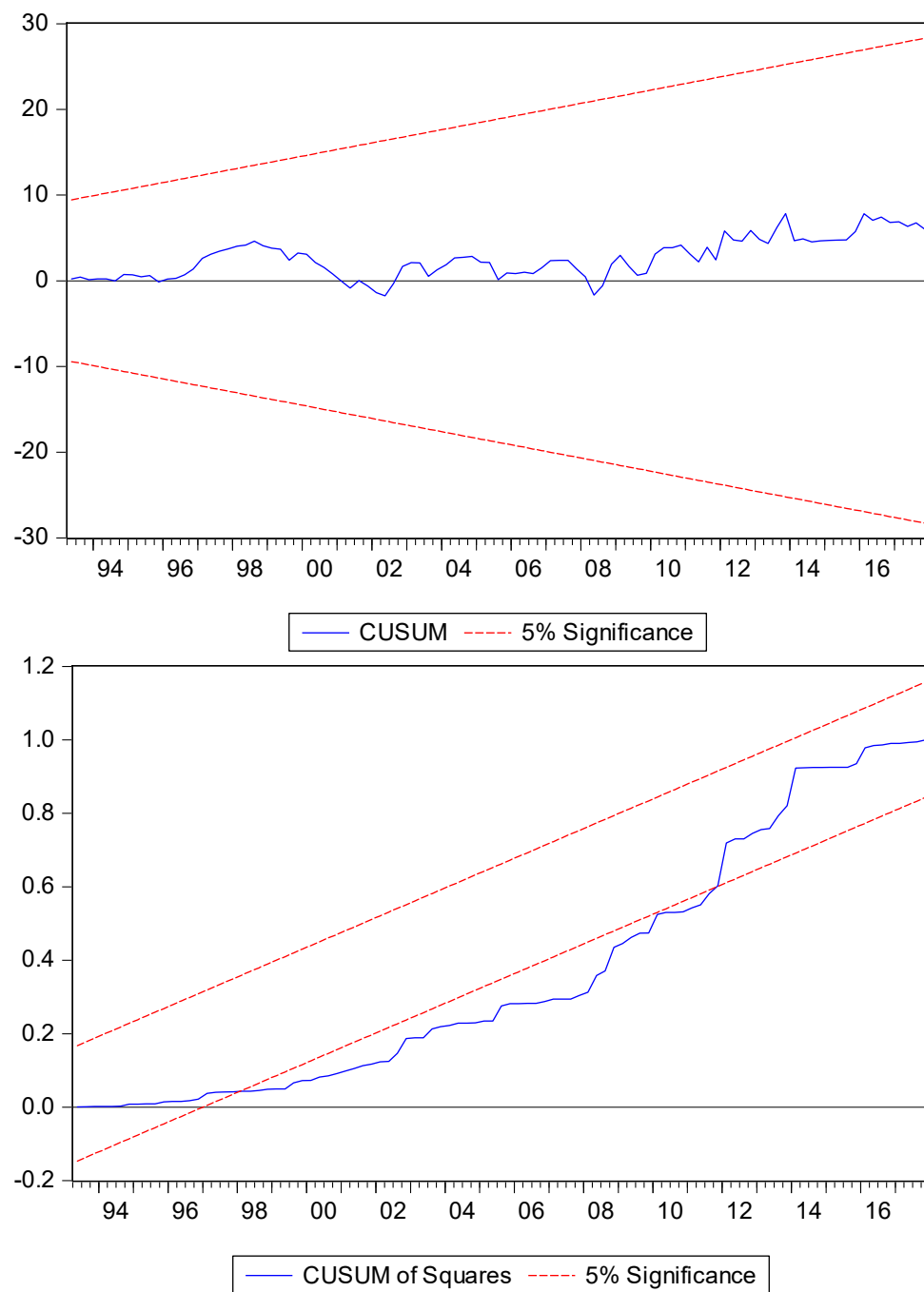
$$ECM_t = LnM_t - \frac{\hat{\rho}_1}{-\hat{\rho}_0} LnY_t - \frac{\hat{\rho}_2}{-\hat{\rho}_0} Lnr - \frac{\hat{\rho}_3}{-\hat{\rho}_0} Ln\left(\frac{P_t}{P_{t-1}}\right) - \frac{\hat{\rho}_4}{-\hat{\rho}_0} LnEX_t - \frac{\hat{\rho}_5}{-\hat{\rho}_0} LnPU_t.$$

We then move back to an error-correction model (2) and replace the linear combination of lagged level variables by ECM_{t-1} and estimate the new specification at the same optimum lags. A significantly negative coefficient obtained for ECM_{t-1} implies that variables are adjusting toward their long-run equilibrium values or cointegrating. The value of the estimated coefficient measures the speed of adjustment. Since the *t*-test is used to judge significance of the coefficient attached to ECM_{t-1}, the test is also known as the t-test for cointegration. Like the F test, Pesaran et al. (2001, p. 303) provided new critical values for this test too. Note also that the value.

Table 2. Full-information estimates of the Linear Autoregressive Distributed Lag (ARDL) Model (2).

Panel A: Short-Run Coefficient Estimates												
Lag Order	0	1	2	3	4	5	6	7	8	9	10	11	12
ΔLnM	-												
ΔLnY	−0.32 [a] (−2.08) *												
ΔLn r	−0.01 (−2.02) *	−0.002 (−0.94)	0.002 (0.52)	0.004 (1.34)	−0.01 (−2.24) *	0.004 (1.57)	−0.004 (−1.32)	−0.003 (−0.60)	0.01 (1.97)				
ΔLn(P_t/P_{t-1})	−0.02 (−4.54) **	0.01 (1.33)	0.01 (2.67) **										
ΔLnLEX	0.11 (2.93) **	−0.04 (−1.24)	0.07 (2.11) *	0.003 (0.09)	−0.08 (−2.81) **	0.05 (2.80) **							
ΔLnPU	0.01 (3.98) **												

Panel B: Long-Run Coefficient Estimates					
Constant	LnY	Ln r	Ln(P_t/P_{t-1})	LnEX	LnPU
131.45 (0.77)	−14.06 (−0.74)	0.39 (0.65)	0.51 (0.79)	0.69 (0.89)	−0.87 (−0.64)

Panel C: Diagnostics					
F [b]	ECM$_{t-1}$	LM [d]	RESET [e]	$\overline{R}^2$	CUSUM (CUSUMQ)
3.46 *	0.01 [c] (0.64)	0.32	23.32 **	0.55	S (UNS)

Notes: a. Numbers inside the parentheses are *t*-ratios. * and ** indicate significance at the 10% and 5% levels, respectively. b. The upper bound critical value of the F-test for cointegration when there are five exogenous variables is 3.35 (3.79) at the 10% (5%) level of significance. These come from Pesaran et al. (2001, Table CI, Case III, p. 300). c. The critical value for significance of ECM$_{t-1}$ is −3.86 (−4.19) at the 10% (5%) level when k = 5. The comparable figures when k = 6 in the nonlinear model are −4.04 and −4.38, respectively. These come from Pesaran et al. (2001, Table CII, Case III, p. 303). d. LM is the Lagrange Multiplier statistic to test for autocorrelation. It is distributed as χ^2 with one degree of freedom. The critical value is 2.70 (3.84) at the 10% (5%) significance level. e. RESET is Ramsey's test for misspecification. It is distributed as χ^2 with one degree of freedom. The critical value is 2.70 (3.84) at the 10% (5%) significance level.

Figure 2. Graphical Presentation of the CUSUM and CUSUMSQ Stability Tests of the Linear Model (2).

The estimates of the nonlinear ARDL model in Table 3 show that introducing nonlinear adjustment of the policy uncertainty measure can change the outcome.

Table 3. Full Information Estimate of the Nonlinear ARDL Model (4) with 0% Threshold.

Panel A: Short-Run Coefficient Estimates												
Lag Order	**0**	**1**	**2**	**3**	**4**	**5**	**6**	**7**	**8**	**9**	**10**	**11**	**12**
ΔLnM	-	−0.20 [a] (−2.93) **	−0.23 (−4.03) **	−0.09 (−1.34)	0.02 (0.38)	−0.19 (−2.26) *							
ΔLnY	−0.39 (2.53) **												
ΔLn r	−0.01 (−5.56) **	−0.002 (−1.13)	0.003 (1.31)	0.01 (2.31) **	−0.003 (−1.34)	0.004 (2.62) **	−0.005 (−2.34) **	−0.004 (−1.01)	0.01 (5.23) **				
ΔLn(Pt/Pt−1)	−0.02 (−5.04) **	0.001 (0.40)	0.003 (0.84)	0.01 (1.26)	−0.003 (−0.65)	−0.01 (−1.51)	−0.01 (−1.29)	0.02 (3.24) **	−0.02 (−4.49) **				
ΔLnLEX	0.08 (3.37) **	−0.02 (−0.63)	0.06 (1.70)	−0.01 (−0.31)	−0.08 (−2.88) **	0.02 (0.78)	−0.01 (−0.24)	−0.01 (−0.52)	0.07 (2.05) *	−0.10 (−4.24) **			
ΔPOS	0.002 (0.32)	−0.01 (−1.94) *	0.01 (1.22)	0.01 (2.35) **	0.004 (1.10)	0.01 (3.17) **	−0.01 (−2.27) *	0.01 (1.45)	0.01 (2.48) **	0.01 (2.35) **	0.01 (2.37) **		
ΔNEG	0.02 (2.60) **												

Panel B: long-run coefficient estimates						
Constant	LnY	Ln r	Ln(Pt/Pt−1)	LnEX	POS	NEG
−4.29 (−0.46)	0.49 (0.52)	−0.28 (−3.31) **	0.13 (5.10) **	1.52 (5.33) **	−0.78 (−3.81) **	0.22 (1.58)

Panel C: Diagnostics							
F [b]	ECM$_{t-1}$ [c]	LM [d]	RESET [e]	$\overline{R}^2$	CUSUM (CUSUMQ)	Wald-L [f]	Wald-S
6.08 **	−0.07 (−2.95)	0.29	2.57	0.67	S (UNS)	9.01 **	58.30 **

Notes: a. Numbers inside the parentheses are *t*-ratios. * and ** indicate significance at the 10% and 5% levels, respectively. b. The upper bound critical value of the F-test for cointegration when there are five exogenous variables is 3.35 (3.79) at the 10% (5%) level of significance. These come from Pesaran et al. (2001, Table CI, Case III, p. 300). c. The critical value for significance of ECM$_{t-1}$ is −3.86 (−4.19) at the 10% (5%) level when k =5. The comparable figures when k = 6 in the nonlinear model are −4.04 and −4.38, respectively. These come from Pesaran et al. (2001, Table CII, Case III, p. 303). d. LM is the Lagrange Multiplier statistic to test for autocorrelation. It is distributed as χ^2 with one degree of freedom. The critical value is 2.70 (3.84) at the 10% (5%) significance level. e. RESET is Ramsey's test for misspecification. It is distributed as χ^2 with one degree of freedom. The critical value is 2.70 (3.84) at the 10% (5%) significance level. f. Both Wald tests are also distributed as χ^2 with one degree of freedom. The critical value is 2.70 (3.84) at the 10% (5%) significance level.

From the short-run results in Panel A we gathered that, again, each variable carries at least one significant coefficient, implying that all variables did have short-run effects. Almost all coefficient estimates obtained for ΔPOS and ΔNEG were positive, supporting the fact that an increase in policy uncertainty raises the demand for money and declines in uncertainty reduces it. However, the short-run effects were asymmetric, not just in terms of their sizes, but also in terms of the adjustment process. Evidence of adjustment asymmetry is borne out by the fact that ΔPOS takes much more lag order than ΔNEG. Strong evidence of short-run asymmetric effects is borne out by the fact that the sum of the coefficients attached to ΔPOS is significantly different than the sum attached to ΔNEG. This is reflected in the significant Wald test that is reported in Panel C as Wald-S, since the test is for short-run cumulative or impact asymmetry. However, only the short-run effects of the POS variable lasts into the long run, since the POS variable carried a significant long-run coefficient in Panel B but the NEG variables did not. This is a clear sign of long-run asymmetric effects, which is also supported by the Wald test reported as Wald-L in Panel C. Furthermore, since the POS variable carries a significantly negative coefficient, it appears that increased policy uncertainty in the U.S.A. induces the public to hold less cash in the long run and more safe financial or real assets so that they can hedge against an uncertain future.

The long-run estimates are meaningful since cointegration is supported by the F-test if not also by the ECM_{t-1} test. Other diagnostic statistics, such as the LM and RESET tests, indicate that there is no evidence of serial correlation and misspecification. Since the nonlinear model is correctly specified and it enjoys a relatively better fit as evidenced by the size of adjusted R^2, the nonlinear model is preferred to the linear model.

Finally, for sensitivity analysis, we excluded changes in the policy uncertainty that were less than 2%. This amounts to changing the threshold level from zero to 2% in generating the partial sum variables as follows:

$$POS_t = \sum_{j=1}^{t} \max(\Delta LnPU_j, 0.02), \ NEG_t = \sum_{j=1}^{t} \min(\Delta LnPU_j, 0.02) \tag{5}$$

The results using the new partial sum variables and nonlinear model (4) are reported in Table 4, and as can be seen, the outcome did not change. The only improvement was support for stability of all coefficient estimates by both CUSUM and CUSUMSQ tests.

Table 4. Full-information estimates of the Nonlinear Model (4) with 2% Thresholds.

Panel A: Short-Run Estimates												
Lag Order	0	1	2	3	4	5	6	7	8	9	10	11	12
ΔLnM	-												
ΔLnY	−0.17 (−1.16) [a]	0.11 (0.66)	0.08 (0.47)	0.32 (1.69)	−0.86 (−4.99) **	0.46 (3.20) **	0.15 (1.06)	−0.15 (−0.97)	0.28 (2.12) *	−0.50 (−4.10) **	0.19 (1.97)		
ΔLn r	−0.01 (−5.34) **	0.005 (2.71) **	0.001 (0.33)	0.01 (2.29) *	−0.005 (−2.02) *	0.001 (0.56)	−0.001 (−0.62)	−0.003 (−0.88) *	0.01 (3.62) **				
$\Delta Ln(P_t/P_{t-1})$	−0.02 (−3.96) **	0.003 (0.88)	0.005 (1.15)	0.01 (2.46) **	−0.01 (−2.58) **	0.0004 (0.10)	−0.01 (−3.44) **	0.02 (4.03) **	−0.01 (−4.42) **				
ΔLnLEX	0.10 (4.06) **	−0.03 (−1.13)	0.01 (0.35)	0.01 (0.36)	−0.08 (−3.28) **	0.04 (1.43)	0.01 (0.45)	−0.06 (−2.17) *	0.07 (2.80) **	−0.14 (−5.15) **	0.07 (3.38) **		
ΔPOS	0.01 (3.09) **	−0.02 (−2.31) **	0.01 (2.59) **	0.01 (2.47) **	0.01 (1.71)	0.01 (1.23)	−0.004 (−0.79)	−0.01 (−1.21)	0.03 (6.24) **	0.01 (2.20) *			
ΔNEG	0.01 (1.67) *	0.01 (1.96) *	0.01 (0.74)	0.0001 (0.01)	−0.003 (−0.54)	0.01 (0.99)	0.004 (0.65)	0.02 (3.61) **	−0.02 (−5.62) **				

Panel B: long-run coefficient estimates						
Constant	LnY	Ln r	$Ln(P_t/P_{t-1})$	LnEX	POS	NEG
1.81 (0.26)	0.12 (0.15)	−0.18 (−3.41) **	0.09 (3.33) **	0.95 (6.73) **	−0.50 (−4.03) **	−0.08 (−0.53)

Panel C: Diagnostics							
F [b]	ECM_{t-1} [c]	LM [d]	RESET [e]	$\overline{R}^2$	CUSUM (CUSUMQ)	Wald-L [f]	Wald-S
3.77 *	−0.09 (−3.08)	0.03	2.36	0.68	S (S)	3.71 *	31.62 **

Notes: a. Numbers inside the parentheses are *t*-ratios. * and ** indicate significance at the 10% and 5% levels, respectively. b. The upper bound critical value of the F-test for cointegration when there are five exogenous variables is 3.35 (3.79) at the 10% (5%) level of significance. These come from Pesaran et al. (2001, Table CI, Case III, p. 300). c. The critical value for significance of ECM_{t-1} is −3.86 (−4.19) at the 10% (5%) level when k =5. The comparable figures when k = 6 in the nonlinear model are −4.04 and −4.38, respectively. These come from Pesaran et al. (2001, Table CII, Case III, p. 303). d. LM is the Lagrange Multiplier statistic to test for autocorrelation. This is distributed as χ^2 with one degree of freedom. The critical value is 2.70 (3.84) at the 10% (5%) significance level. e. RESET is Ramsey's test for misspecification. This is distributed as χ^2 with one degree of freedom. The critical value is 2.70 (3.84) at the 10% (5%) significance level. f. Both Wald tests are also distributed as χ^2 with one degree of freedom. The critical value is 2.70 (3.84) at the 10% (5%) significance level.

4. Concluding Remarks

In the early 1980s, when the Fed missed its inflation target, Friedman (1984) attributed it to volatility of monetary growth rate. Since then, many studies have tried to assess the impact of alternative measures of uncertainty on the velocity of the money or on the demand for cash balances. The literature has advanced on two grounds. First, relatively more comprehensive measure of uncertainty is being used to assess public's demand for money in response to an uncertainty measure. The new uncertainty measure includes any factor that contributed to an uncertain economic and political environment. Second, in some countries such as Australia and Korea, previous research has discovered that public respond to changes in uncertainty measure in an asymmetric manner.

In this paper, we estimated the demand for money in the U.S.A. by including a new measure of uncertainty known as policy uncertainty that was constructed by Policy Uncertainty Group (see the Appendix A) by searching popular newspapers in the U.S.A. Bahmani-Oskooee et al. (2016), who used the new measure and assessed its impact on the demand for money in the U.S.A., assumed that the effects were symmetric. In this paper, we deviated from that assumption and argued for asymmetric effects. Since we used updated data, we first estimated the demand for money by holding symmetric assumption. This amounts to applying the linear ARDL approach from Pesaran et al. (2001). Then, we changed the symmetry assumption to asymmetry assumption and used the nonlinear ARDL approach by Shin et al.

Our findings are best summarized by saying that when the linear ARDL model was estimated, policy uncertainty had short-run effects but not long-run effects. However, estimates of the nonlinear ARDL model revealed that changes in policy uncertainty have both short-run and long-run effects on the demand for money in the U.S.A. Furthermore, both the short-run and long-run effects were asymmetric.[9] In the long run, we found that while increased policy uncertainty had adverse effect on the demand for cash, decreased policy uncertainty has no long-run effects. Increased uncertainty makes people more cautious about the future; however, they do not simply adjust their portfolio to decreased uncertainty, perhaps because they are used to some degree of uncertainty in their environment, or because they expect more uncertainty in the future.

Author Contributions: M.M.-N. collected the data, carried out the entire estimation, and tabulated the results, M.B.-O. wrote the paper.

Appendix A

Appendix A.1 Data Definition and Sources

Quarterly data over the period 1985I–2017IV were used to carry out the estimation. The main restriction for using data prior to 1985 was unavailability of data on policy uncertainty. Data were collected from the following sources:

(a)　International Financial Statistics (IFS) of International Monetary Fund (IMF).
(b)　Economic Policy Uncertainty Group: http://www.policyuncertainty.com/us_monthly.html
(c)　Federal Reserve Bank of St. Louis (FRED)

Appendix A.2 Variables

$M2$ = Money supply measured by real M2. Data come from Source (c).
Y = Real GDP. Data come from Source (c).
r = Interest rate. Interest rate on 3-Month Treasury Bill from Source (c).

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[9]　These finding for the U.S.A. are similar to those of Bahmani-Oskooee and Maki-Nayeri (2018a, 2018b) for Korea and Australia respectively.

P = Price level used to measure the inflation rate is GDP deflator. Data come from Source (c).

EX = Index of nominal effective exchange rate of the U.S dollar. A decline reflects a depreciation of U.S dollar. Data come from Source (a).

PU = Policy uncertainty. Data come from Source (b).

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