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A Note on the Relationship Between the Cyclicality of Markups and Fiscal Policy*

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1. Introduction

The role of endogenously varying markups in the transmission mechanism of fiscal policy has been a matter of interest in the recent macroeconomic literature, e.g. [4], [11], [26]. However, empirical evidence on the cyclical behavior of markups following a government spending shock is mixed: [20] find evidence of counter-cyclical markups in the US; [1] find mixed results for other OECD countries; [22] obtain either acyclical or pro-cyclical markups in US industries.

We demonstrate that, conditional on the fiscal shock, the cyclical behavior of markups is related to the cyclical behavior of government spending. Under plausible parameter assumptions, pro-cyclical spending implies less counter-cyclical markups. Evidence on a dataset 1970-2007 for 13 OECD countries, for which we test and compare the cyclical properties of markups and total final government spending, confirms a weak version of this theoretical hypothesis.

2. Composition effects

Let us assume that both households and government have CES preferences over a continuum of differentiated goods with mass equals to one. The private, $c(j)$, and public, $g(j)$, demand functions for good $j \in [0,1]$ are given by

$$c(j) = \left(\frac{p(j)}{P_C}\right)^{-\sigma} C \ , \quad g(j) = \left(\frac{p(j)}{P_G}\right)^{-\varepsilon} G \ , \quad (1)$$

where $p(j)$ is the price, C (G) is real aggregate private (public) consumption, P_C (P_G) is the appropriate private (public) price index. The private and the public sector only differ in their degree of elasticity of substitution amongst varieties: σ, $\varepsilon > 0$.

If each good is produced under monopolistic competition, then firms face market demands $y(j) = c(j) + g(j)$ with elasticity given by

$$\eta(j) = \sigma(1-\gamma(j)) + \varepsilon\gamma(j) \ , \quad (2)$$

where $\gamma(j) \equiv g(j)/y(j)$ is the share of public demand on output of good j. In this model the markup is given by $\mu(j) \equiv p(j)/m(j) = \eta(j)/(\eta(j)-1)$, where $m(j)$ stands for the marginal cost. In a symmetric equilibrium we have that γ is a function of aggregate output $Y = C + G$. Therefore, as already noticed by [13][1] or [8][2], the composition of aggregate demand affects the average markup, as long as $\sigma \neq \varepsilon$. Thus the cyclical response of markups is related to the cyclical behavior of government spending. We can see this by differentiating the average markup with respect to Y:

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[1] See op. cit. section 4.4.
[2] Using investment instead of government expenditures.

$$\operatorname{sign}\left(\frac{\partial\mu}{\partial Y}\right)=\operatorname{sign}\left(-\frac{\eta'}{(\eta-1)^2}\right)=\operatorname{sign}\left((\sigma-\varepsilon)\frac{\partial\gamma}{\partial Y}\right) \tag{3}$$

If $\sigma > \varepsilon$ the markup moves relative to output in the same direction as the cyclical response of the public expenditure share to output. Hence, the cyclical response of the markup should mimic the cyclical behavior observed for the government share in final demand.

We should expect σ to be greater than ε, as most of the government spending is non-tradable (e.g. defense, health, education), depends on a small number of suppliers, and is often bound by long-run contracts that tend to be stricter on quantities than they are on prices[3]. In fact, using the US input-output tables (mid 2000's), we find a positive rank correlation (about 0.3) between the share of purchases of sector C75 (public administrations and defense; compulsory social security) from other sectors and the markup measure of these industries in both [5], and [19]. Therefore, we tend to observe, at least for the US, larger markups, i.e. smaller elasticities faced by firms in sectors that supply a larger share of its sales to the public sector.

If $\sigma > \varepsilon$, then we should observe counter(pro)-cyclical average markups - following an increase in government spending - in economies where public spending is also counter(pro)-cyclical.

3. Pro-cyclical fiscal policy

Most models assume that G is determined exogenously, either in a deterministic or in a stochastic way. Therefore, the share of government expenditure in output is counter-cyclical in that case and thereby inducing counter-cyclical markups[4]. Yet, empirical evidence shows that governments often respond pro-cyclically to economic conditions, even in OECD countries - see [17] and [7]. Theorists explain this phenomenon by pressure from multiple power blocs for additional spending when economic booms generate budget surpluses - see [2], [24], or [25].

Hence, we cannot *a priori* suppose $\gamma(Y)$ to be decreasing. To test the degree of pro-cyclicality of spending, numerous papers, e.g. [17], have estimated an equation like

$$\Delta \ln G_t = \alpha + \beta \Delta \ln Y_t + u_t \ . \tag{4}$$

If $\beta < 0$, spending falls during an upswing in the cycle, and we call spending counter-cyclical. Instead, if $\beta > 0$, spending is pro-cyclical. It is weakly pro-cyclical if the rise in spending is less than proportional to the change in output ($0 < \beta < 1$), and strongly pro-cyclical if the response is more than proportional ($\beta > 1$)[5]. Assuming $\sigma > \varepsilon$ and using (3) the relation between average markups and the cyclicality of spending is positive only when spending is strongly pro-cyclical.

4. Evidence

We revisit evidence on the degree of cyclicality of public spending and then compare it to the cyclical pattern of average markups. This pattern is taken from [1] who study the cyclicality of average markups following a fiscal shock using a VAR approach for a sample of 14 OECD countries in the 1970-2007 period. They find significantly negative reactions of average markups to a government spending shock for all countries with the exception of the US and Sweden[6]. The column 'markups' on Table 1 shows the sign of the ratio of the impact responses of output and the average markup to a government spending shock.

We use the same dataset as [1] to test the cyclical response of spending in (4). This test should account for the endogeneity of output to changes in the budget. VAR studies inspired by the seminal work of [3] show strong evidence that fiscal

[3] See [22], Table 1, for the top 10 industries with the largest share of shipments to the US government.

[4] Automatic stabilizers bolster this counter-cyclical response. The most important category of built-in stabilizers is unemployment insurance. As this is a small fraction of overall spending, the end-effect is limited - see [9].

[5] Notice that $\gamma'(Y) = \frac{\gamma}{Y}\left(\frac{\partial G}{\partial Y}\frac{Y}{G}-1\right)$ and β provides the corrected estimate for $\frac{\partial G}{\partial Y}\frac{Y}{G}$.

[6] The response is not significant within the 95 per cent bands.

policy does have real economic effects. Instrumental variable (IV) estimators of (4) applied to a panel of high-income countries show mixed evidence, however - see [15] and [14]. Nonetheless, the instruments employed in these papers, albeit exogenous, turn out to be irrelevant or weak and so the estimate of β is biased and its standard error understated. A first remedy is to employ an extended set of instruments that is arguably stronger than in previous studies. We follow [15] and use a real external shock as a first instrument: this shock is the product of a trade-weighted average of real GDP growth of the country's trading partners and the average export share in GDP. Following the same rationale, we also include the growth rate of a reference economy[7]. The increase in exports led by higher growth abroad has no direct effect on government spending, but affects the domestic business cycle. Both instruments may not be exogenous for large economies that move, rather than endure, output in trading partners. In addition to these commonly employed instruments, we add three new instruments. First, we include commodity imports as given by the product of the annual change in the Commodity Research Bureau price index and the average import share of commodities in GDP. Higher raw-materials costs dampen growth, yet they should not lead to changes in government spending[8]. Second, changes in the labor force shift the production frontier, so it is correlated with (long-term) economic growth, but is also orthogonal to spending. Finally, we propose using some financial flows as instruments. Remittances and private aid flows (both grants and FDI) measure financial flows out of the country[9] and, although they do not affect the domestic business cycle, they arguably follow it[10].

A second remedy is to test instruments for their weakness and to employ a battery of IV estimators that are robust to weak instruments. We compute the heteroskedasticity- and autocorrelation-consistent (HAC) robust first-stage Lagrange Multiplier (LM) test for weak instruments of [16], henceforth KP. We reject the null of weak instruments if the KP-statistic is larger than the Stock-Yogo critical values. If instruments are weak, we correct the bias in the coefficient estimates and size distortion in hypothesis tests by estimating (4) with a Continuous Updating Estimator (CUE). This estimator is a GMM-like HAC robust version of a Limited Information Maximum Likelihood (LIML) estimator and provides more reliable point estimates under weak instruments than either GMM or LIML - see [12]. Alternatively, we apply the conditional Likelihood Ratio (LR) test using the pivotal statistics of [21] that provides a confidence interval with the correct size[11]. We additionally show two Fuller-k estimators that also provide good robust estimates for in this case - see [10]. The KP and J tests reported in Table 1 show that the set of instruments is not weak and valid for all countries. The Fuller-k estimators are therefore approximately unbiased, and inference with the IV estimators correct. Only in Finland and Sweden the set of instruments is probably weak and we should look at the conditional LR test. The resulting CUE of β is significant in nearly all countries (Belgium is the exception) and its sign is also consistent across estimators although the Fuller-k estimates are not always significant.

A comparison of the sign of β with the cyclicality indicator implicit in [1] confirms our theoretical hypothesis: both the signs of β and the cyclicality indicator are equal (with the exception of Sweden). However, we cannot provide support to a strong version of our model: average markups should be pro-cyclical if public spending is also strongly pro-cyclical ($\beta > 1$). We find that spending is close to being strongly pro-cyclical

[7] Germany for EU countries, US for non-EU countries and Germany.

[8] This instrument may be less adequate for some commodity exporters (Australia, Canada) as windfall tax revenues may fuel additional spending.

[9] Data on remittances comes from World Bank estimates, as described in [23]. Bilateral aid data is available from the OECD DAC database.

[10] Immigrants' remittances have been shown to be pro-cyclical with respect to GDP in the donor country - see [18]. Similarly, development aid is positively correlated with GDP: good economic times increase generosity towards developing countries, whereas such aid is cut during economic downturns - see [6].

[11] This LR test is only valid for a single endogenous regressor, which is the case in (4).

in Australia, Belgium[12], France, and the Netherlands, but since point estimates of β are positive at 95 per cent (according to the conditional LR interval), spending can only be considered to be weakly pro-cyclical. A borderline case is Finland: as instruments are rather weak, we can only use the conditional LR interval which shows that the data is consistent with either a counter- or pro-cyclical response of spending. All other countries display a counter-cyclical markup and also pursue a counter-cyclical policy in spending with all CUE point estimates rejecting a positive response. Even when instruments are marginally weak in Canada and the UK, the Moreira confidence interval mostly covers negative values. Evidence is weaker for Sweden: government spending looks clearly counter-cyclical, but there is so much uncertainty on the responses to a fiscal shock that we cannot rule out the possibility of a counter-cyclical markup either. In a nutshell: the cyclical reaction of government spending to output matches that of the cyclical response of average markups, which is consistent at least with a weak version of our theoretical prediction.

5. Conclusion

In this paper, we demonstrate that, in a simple endogenous-markup model, the cyclical behavior of average markups is related to the cyclical behavior of government spending. Under plausible parameter assumptions, pro-cyclical spending results in less counter-cyclical mark-ups. We use IV estimators to test cyclicality of government spending in a dataset of 13 OECD countries over the 1970-2007 period and compare the results to markup cyclicality implicit in [1]. We confirm a weak version of our theoretical hypothesis.

Our paper supplies an explanation that reconciles the variety of empirical findings on the cyclical behavior of markups with the theoretical literature on endogenously varying markups and fiscal policy.

[12] None of the estimators of β is significant, and so we cannot reject β to be smaller than 0 nor larger than 1.

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Table 1
Ciclality of Markups and Fiscal Policy.

Country s	Markups	IV estimation of β					
		KP	J	CUE	Fuller(2)	Fuller(4)	Conditional LR
Australia	+	22.02	0.73	0.28**	0.39**	0.37**	[-0.02,0.92]
				(2.18)	(2.25)	(2.30)	0.03/0.97/0.03
Belgium	+	32.66	0.51	0.22	0.15	0.14	[-0.49,0.82]
				(0.91)	(0.48)	(0.49)	0.60/0.02/0.02
Canada	-	12.79	0.80	-0.34**	-0.19	-0.16	[-0.93,0.36]
				(-2.05)	(-0.96)	(-0.90)	0.09/0.25/0.00
Denmark	-	17.97	0.63	-0.57**	-0.05	-0.01	[-3.55,1.18]
				(-2.64)	(-0.14)	(-0.02)	0.78/0.39/0.06
Finland	- (a)	7.75	0.54	-0.26	0.16	0.18	[-0.81,0.80]
				(-1.44)	(0.31)	(0.39)	0.63/0.68/0.02
France	+	13.61	0.80	0.47***	0.36	0.26	$(-\infty,+\infty)$
				(4.03)	(0.76)	(0.88)	0.61/0.70/0.84
Germany	-	103.43	0.81	-0.35***	-0.23	-0.10	[-3.10,1.35]
				(-3.30)	(-0.47)	(-0.23)	0.62/0.31/0.12
Italy	-	46.16	0.67	-0.37**	-0.35	-0.30	[-1.44,0.38]
				(-1.99)	(-0.49)	(-0.46)	0.40/0.20/0.00
Japan	-	26.54	0.71	-0.84**	-0.07	-0.09	[-0.67,0.68]
				(-2.29)	(-0.18)	(-0.24)	0.89/0.45/0.01
Netherlands	+ (a)	38.78	0.62	0.22***	0.06	0.06	[-0.49,0.55]
				(2.99)	(0.25)	(0.29)	0.84/0.58/0.00
Sweden	+ (a,b)	2.53	0.61	-0.67***	0.38*	-0.33*	[-1.31,0.03]
				(-4.40)	(1.85)	(-1.79)	0.06/0.03/0.00
UK	-	11.51	0.72	-0.48***	-0.50**	-0.49**	[-1.78,0.70]
				(-3.55)	(-2.22)	(-2.31)	0.26/0.13/0.02
USA	- (b)	92.10	0.59	-1.46***	-1.97	-1.56	[-12.99,-0.68]
				(-2.83)	(-0.53)	(-0.56)	0.01/0.00/0.00

NOTES: 'Markups' stands for $\text{sign}\left(\frac{\Delta \ln Y_{s1}\left(\varepsilon_{s1}^{G}\right)}{\ln \mu_{s1}\left(\varepsilon_{s1}^{G}\right)}\right)$ in [1], where ε_{s1}^{G} represents an unexpected shock to G in country s occurring at time $t = 1$; 'KP' represents [16] Wald rk F statistic for weak instruments: Stock-Yogo critical values for single endogenous regressor for a 5(10) per cent maximal IV relative bias 13.91(9.08) and 5(10) per cent maximal Fuller relative bias 9.61(7.90); "J" is the J test for validity of instruments; 'CUE,' 'Fuller(2),' and 'Fuller(4)' β estimates, significance, and t-statistics (in brackets); 'Conditional LR' 95 per cent confidence interval for β (the three p-value for null hypotheses $\beta = 0$, $\beta > 0$, and $\beta = 1$); (a) a zero response of Y to an impulse on G is within the confidence bands; (b) a zero response of μ to an impulse on G is within the confidence bands; ***/**/* significance at 1/5/10 per cent.

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