A Simple Approach to Testing the Potency of Government Purchases to Stimulate Aggregate Demand

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Abstract: The paper proposes a new approach to testing the potency of government purchases to stimulate the economy by testing a set of conditions implied by the Ricardian Equivalence (RE) proposition that a typical household incorporates the government's budget constraint into its own. These conditions are as follows: (1) private consumption, income, and government purchases form a "levels relationship"; and (2) considering consumption as the dependent variable, the coefficients of income and of government purchases are 1 and -1. The last restriction is also implied by the hypothesis that consumption and government purchases are perfect substitutes, however, so the proposed approach cannot distinguish between the perfect substitutability and the RE hypotheses. This restriction is thus referred to in the paper as the hypothesis of direct or *ex ante* full crowding out. If it holds, then the multiplier of government purchases is zero. Using US quarterly data, 1947.1-2012.1, the results suggest that a "levels relationship" exists and that the coefficient of government purchases is about -0.4 and significantly below -1, thus leading to the conclusion that government purchases stimulate aggregate demand and output.

Keywords: Ricardian equivalence, government purchases, *ex ante* crowding out.

1. INTRODUCTION

Recently, thanks to the global economic crisis, there has been an upsurge in the debate on the size (and the sign!) of the fiscal multipliers (Ramey 2011; Auerbach and Gorodnichenko 2012; Ilzetzki, Mendoza, and Végh 2013). This paper proposes an approach to testing the effect of government purchases on private consumption, and hence on aggregate demand. This effect can occur through three channels: (1) Ricardian Equivalence (RE); (2) indirect or *ex post* crowding out (through the interest rate); and (3) substitutability (or complementarity) between government purchases and private consumption. The model proposed here is an implication of RE.

In a continuous time setting, Romer (2006, section 11.2) shows that a key result of RE is that it does not matter for the economy whether a given stream of government purchases is financed by taxes or by bonds; it is the level of that stream that matters. This result follows if it is assumed that (1) taxes do not enter directly a typical household's utility function, and (2) the household incorporates the government's budget constraint into its own, and this "combined" budget constraint involves government purchases, but not taxes or government debt. Thus, for example, a bond-financed permanent increase in government purchases will cause rational consumers to reduce their consumption one for one, so that they can increase their saving and pay the expected higher taxes in the future. That is, RE implies that government purchases fully crowd out private consumption, so they do not affect aggregate demand, implying a zero government-purchases multiplier. This example illustrates the direct or *ex ante* crowding out effect implied by RE. More generally, as Miller (1982:420) notes, *ex ante* crowding out occurs when "changes in fiscal policy cause behavioral adjustments that crowd out private demand *before* other system variables are affected" (emphasis in the original).

Aschauer (1985) and others test the RE hypothesis by assuming a great deal of rationality on the part of a typical household, e.g., intertemporal utility maximization, RE, and rational forecasting of future consumption and government purchases. Their models give rise to a large number of nonlinear cross-equation restrictions among a large number of parameters. Such a "heavy" structure requires a large sample, however, so that the tests can have reasonable power.

In contrast, the model proposed here exploits only the "combined" budget constraint described above. The estimating equation is analogous to that of Hakkio and Rush (1991), who developed a well-known test for budget deficit sustainability based on a cointegrating regression derived from the government's budget constraint. If RE holds, the model must satisfy the following conditions: (1) consumption, income, and government purchases form a "levels relationship"; and (2) considering consumption as the dependent variable, the coefficients of income and of government purchases are 1 and -1. The last restriction is also implied by the perfect substitutability hypothesis, however, which also gives rise to *ex ante* full crowding

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out, since it assumes that government purchases substitute for private consumption one for one (Feldstein 1982:9). Thus, in the present paper this restriction is referred to as the *ex ante* full crowding out hypothesis. It follows that the proposed approach cannot distinguish between the hypotheses of perfect substitutability and RE.

In addition to the identification problem just described, note that the budget constraint is only a necessary, but not a sufficient, condition for RE (Haug 1991:97). Thus, the proposed approach, which exploits only the "combined" budget constraint, cannot be used to test RE. It rather tests the hypothesis that government purchases fully crowd out private consumption, thus testing the potency of government purchases to stimulate aggregate demand and output. From a policy perspective, this is the question that matters. Since *ex ante* full crowding out implies that the coefficient of government purchases is -1, it follows that an estimate of this coefficient that is significantly lower than -1 (in absolute value) leads to a rejection of the *ex ante* full crowding out. The model proposed here resembles the traditional consumption functions that have been used to test RE without using cointegration analysis (Feldstein 1982; Kormendi 1983). Its advantages over previous studies are (1) simplicity, (2) use of modern cointegration methods, and (3) test power (on *a priori* grounds), since only two parameters are estimated and tested.

Note that because of the important policy implications, new models have been developed recently, which relax some of the strong assumptions made by Aschauer (1985), such as the absence of liquidity constraints. Thus, new evidence has emerged, especially from international data (Khalid 1996; Giorgioni and Holden 2003; Reitschuler and Cuaresma 2004; Nieh and Ho 2006; Ilzetzki *et al*. 2013). The empirical evidence is mixed, however. This suggests that a more powerful testing approach is needed to answer the important policy question whether government purchases can stimulate aggregate demand. The present paper is an attempt toward that direction. Section 2 develops the new approach and section 3 applies it to United States (US) data. Section 4 concludes.

2. THE MODEL

In discrete time, the period-by-period budget constraint of a typical household in real terms is given by [see, e.g., Aschauer (1985), Equation (2)]

$$A_{t+1} = (1 + i)(A_t + H_t + tr_t - tx_t - C_t), \qquad (1)$$

where A_t = the household's beginning-of-period non-human wealth, including government debt and income flows that are automatically capitalized, e.g., interest income from bank deposits or private debt held in period $t - 1$, but excluding interest income from holding government debt; i = interest rate, which is assumed to be constant (for analytical tractability), as in Aschauer (1985); H_t = labor income; tr_t = government transfers, including interest on the government debt held in period $t - 1$; tx_t = taxes; and C_t = total consumption expenditure on goods and services.

The government's period-by-period budget constraint (also in real terms) is

$$B_{t+1} = (1 + i)(B_t + G_t + tr_t - tx_t), \qquad (2)$$

where B_t = beginning-of-period government debt and G_t = government purchases of goods and services. Subtracting (2) from (1) yields the "combined" budget constraint

$$a_{t+1} = (1 + i)(a_t + H_t - G_t - C_t), \qquad (3)$$

where $a_t = A_t - B_t$ is the household's wealth net of government debt. Solving the difference equation (3) forward and imposing the solvency condition

$$\lim_{T\to\infty}[a_{t+T} / (1+i)^T] = 0 \qquad (4)$$

yields

$$a_t = \sum\nolimits_{s=0}^{T-1} \beta^s (C_{t+s} + G_{t+s} - H_{t+s}), \qquad (5)$$

where $\beta = 1/(1+i)$, $0 < \beta < 1$. Equation (5) coincides with Aschauer's (1985) Equation (6) and is the discrete-time analog of Romer's (2006) Equation (11.11).

Following Hakkio and Rush (1991) in the derivation of their Equation (6), set $C_{t+s} = \Delta C_{t+s} + C_{t+s-1}$, $G_{t+s} = \Delta G_{t+s} + G_{t+s-1}$, and $H_{t+s} = \Delta H_{t+s} + H_{t+s-1}$ in Equation (5), where Δ is the first-difference operator. After some algebraic manipulation, the following equation emerges:[1]

$$Y_t - C_t - G_t = \frac{1}{\beta}\sum\nolimits_{h=1}^{\infty} \beta^{h-1} (\Delta C_{t+h} + \Delta G_{t+h} - \Delta H_{t+h}), \qquad (6)$$

where $Y_t = H_t + i(a_{t+1} + H_t - C_t - G_t)$ is labor and interest income before taxes and transfers, and where

[1]Equation (6) is derived in an appendix, which is available from the author upon request.

capitalization of interest is assumed to occur at the end of period t. Assuming that C_t, G_t, and H_t are each a random walk with drift and substituting into Equation (6) $\Delta C_{t+h} = \alpha_1 + \varepsilon_{1,t+h}$, $\Delta G_{t+h} = \alpha_2 + \varepsilon_{2,t+h}$, and $\Delta H_{t+h} = \alpha_3 + \varepsilon_{3,t+h}$, where ε_{1t}, ε_{2t}, and ε_{3t} are white-noise processes, there emerges the following regression equation:

$$C_t = a + b_1 Y_t + b_2 G_t + \varepsilon_t, \tag{7}$$

where $a = (\alpha_3 - \alpha_1 - \alpha_2)/[\beta(1-\beta)]$ and $\varepsilon_t = \beta^{-1}\Sigma_{h=1}^{\infty}\beta^{h-1}(\varepsilon_{3,t+h} - \varepsilon_{1,t+h} - \varepsilon_{2,t+h})$. The model implies the following joint hypothesis: (1) C_t, Y_t, and G_t are cointegrated; and (2) the coefficient restrictions $b_1 = 1$ and $b_2 = -1$ must hold.

Note that whenever there is evidence that the variables C_t, G_t, and Y_t are I(0), then, instead of a "cointegrating relation," Equation (7) should be referred to as a "levels relationship" (Pesaran *et al*. 2001). In this case, we can still proceed as before by assuming, for example, that C_t, G_t, and H_t are AR(1) processes: $C_t = \alpha_1 + \rho_1 C_{t-1} + \varepsilon_{1t}$, $G_t = \alpha_2 + \rho_2 G_{t-1} + \varepsilon_{2t}$, and $H_t = \alpha_3 + \rho_3 H_{t-1} + \varepsilon_{3t}$, where $|\rho_j| < 1$, $j = 1, 2, 3$. Adding $C_{t-1} - C_{t-1}$ to the first, $G_{t-1} - G_{t-1}$ to the second, and $H_{t-1} - H_{t-1}$ to the third of these three equations yields $C_t = \alpha_1 + C_{t-1} + \varepsilon_{1t}^*$, $G_t = \alpha_2 + G_{t-1} + \varepsilon_{2t}^*$, and $H_t = \alpha_3 + H_{t-1} + \varepsilon_{3t}^*$, where $\varepsilon_{1t}^* = \varepsilon_{1t} + (\rho_1 - 1)C_{t-1}$, $\varepsilon_{2t}^* = \varepsilon_{2t} + (\rho_2 - 1)G_{t-1}$, and $\varepsilon_{3t}^* = \varepsilon_{3t} + (\rho_3 - 1)H_{t-1}$. The only effect on Equation (7) is that its error term is now defined as $\varepsilon_t = \beta^{-1}\Sigma_{h=1}^{\infty}\beta^{h-1}[\varepsilon_{3,t+h} - \varepsilon_{1,t+h} - \varepsilon_{2,t+h} + (\rho_3 - 1)H_{t-1} - (\rho_1 - 1) C_{t-1} - (\rho_2 - 1)G_{t-1}]$. This is a stationary variable, as ε_{1t}, ε_{2t}, ε_{3t}, C_t, G_t, and H_t are all assumed to be I(0), so Equation (7) can be estimated by standard econometric methods.

3. APPLICATION TO U.S. DATA

In this section, the above joint hypothesis is tested by using US quarterly data, 1947.1-2012.1, obtained from the US Department of Commerce: Bureau of Economic Analysis, National Income and Product Accounts. The empirical definitions of the variables are as follows (all are expressed in constant 2005 dollars and are seasonally adjusted): Y_t = real per capita net national product; C_t = real per capita personal consumption expenditure on goods and services; and G_t = real per capita Federal, State, and Local government consumption expenditure on goods and services plus gross government investment.

Figure **1** shows that the series C_t and Y_t share a common trend throughout the sample period, but the gap between this trend and that of G_t started to increase in the late 1960's. This fact is taken into account in the cointegration analysis below by including a time trend in the cointegration space.

Begin by testing for unit roots and cointegration. Four unit-root tests are used: (1) the Phillips-Perron test (*PP*); (2) the "point-optimal test" of Elliott, Rothenberg, and Stock (*ERS*, 1996); (3) the *KPSS* test of Kwiatkowski *et al*. (1992); and (4) the Lee and Strazicich (*LS*, 2003, 2004) test, which allows for one or two structural breaks. As cointegration tests, the following three are used: (1) Johansen's trace test; (2) the Gregory and Hansen (*GH*, 1996a, 1996b) tests, which allow endogenously for a structural break; and (3) the Pesaran *et al*. (2001) "bounds test." Table **1** reports the results. Its notes provide more details about these tests.

At the 5% level, the tests of Table **1** suggest that the variables C_t, Y_t, and G_t are all I(1) and form one cointegrating relation. At the 10% level, the Gregory-Hansen test provides weak evidence of a level shift in 2001:2, as the minimum t-statistic is -4.70, whereas the

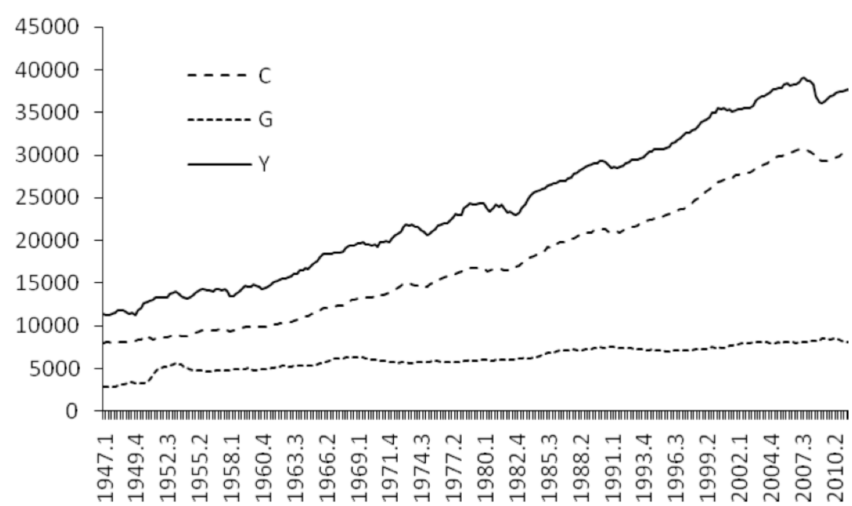

Figure 1: The evolution of the series *Y*, *C*, and *G* from 1947.1 to 2012.1.

Table 1: Unit-Root and Cointegration Tests

Part A. Unit-root tests											
Test Series	PP_μ	PP_τ	ERS_μ	ERS_τ	$KPSS_\mu$	$KPSS_\tau$	***LS* one crash**	***LS* two crashes**	***LS* one break**	***LS* two breaks**	**I(0) or I(1)?**
C_t	1.26	-2.29	604.2	37.3	4.38^{***}	0.84^{***}	-1.65	-1.73	-3.12	-4.42	I(1)
Y_t	-0.03	-2.83	574.9	10.4	4.41^{***}	0.47^{***}	-2.65	-2.77	-4.22^{*} (1997:4)	-5.17	I(1)
G_t	-2.30	-3.07	109.6	10.6	4.01^{***}	0.15^{**}	-3.27^{*} (1969:4)	-3.37	-3.96	-4.71	I(1)
ΔC_t	-12.0^{***}	-12.2^{***}	0.38^{***}	1.27^{***}	0.40^{*}	0.10	-6.41^{***}	-7.24^{***}	-6.84^{***}	-7.31^{***}	I(0)
ΔY_t	-11.3^{***}	-11.3^{***}	0.13^{***}	0.39^{***}	0.07	0.05	-7.89^{***}	-8.01^{***}	-7.82^{***}	-8.41^{***}	I(0)
ΔG_t	-10.5^{***}	-10.6^{***}	0.39^{***}	1.31^{***}	0.20	0.06	-7.11^{***}	-7.31^{***}	-7.20^{***}	-7.31^{***}	I(0)

Part B. Cointegration tests (r = cointegration rank)					
Johansen's trace test	*GH* (*C*)	*GH* (*C* | *T*)	*GH* (Full break)	Bounds test	r = ?
r = 0 vs. r = 1: 58.44^{***} r = 1 vs. r = 2: 27.29^{*} r = 2 vs. r = 3: 7.30	r = 0 vs. r = 1: -4.70^{*} (2001:2)	r = 0 vs. r = 1: -3.78	r = 0 vs. r = 1: -5.09	r = 0 vs. r = 1: 7.49^{***}	1 (at the 5% level)

Notes: (1) ***, **, * indicate significance at the 1%, 5%, and 10% level; (2) in the *PP* and *ERS* tests, the subscripts μ and τ indicate, respectively, "intercept-but-no-trend" and "intercept-plus-trend," whereas in the *KPSS* tests they indicate level and trend stationarity; (3) in every test, the lag length (l) was set equal to 5, based on the formula l = integer[$4(T/100)^{1/4}$] (Kwiatkowski *et al*. 1992:169); here, T = 261; (4) in the *LS* tests, the possible break dates of the levels are given in parentheses underneath the values of the test statistic, provided that these values are significant at the 10-percent level or lower; in the cases of the first differences, no breaks are considered; (5) in Part B, Johansen's trace test is produced by assuming four lags, a level shift, and a time trend in the cointegrating relation, in accordance with Figure **1** and the *GH* (*C*) test, which is significant at the 10% level; (6) *GH* (*C*), *GH* (*C* | *T*), and *GH* (Full break) are Gregory and Hansen's (1996a, 1996b) "level shift," "level shift with trend," and "full break" models, where the maximum lag length was set to 16; (7) in the "bounds test," the maximum lag length was set to 4, and insignificant lags were dropped one at a time; standard errors are robust to heteroscedasticity and autocorrelation; critical values were obtained from Table CI(iii) Case III of Pesaran *et al*. (2001:300); (8) The *ERS* test was implemented using the econometric program *Eviews* 7.0, whereas all the other tests were carried out using *RATS* 7.0 and *CATS in RATS* 2.0.

10% critical value is -4.69; see Gregory and Hansen (1996a, Table **1**) for m = 2. A possible explanation might be that 2001:2 is close to the extraordinary events of 9/11. The ratio G_t/Y_t started to rise slightly in 2001:1 and continued to rise even after the end of the 2001 recession (November 2001), until 2003:2. This result is taken into account below in the cointegration analysis by including in the cointegration space a dummy, denoted as $D2001_t$, which takes on the value of 1 from 2001:2 onward and the value of 0 otherwise. That is, the estimating equation is

$$C_t = a + b_1 Y_t + b_2 G_t + c_1 D2001_t + c_2 t + \varepsilon_t, \tag{8}$$

and the restrictions to be tested are as follows: (1) cointegration; and (2) b_1 = 1 and b_2 = -1. Two methods are used with four lags in each: (1) Johansen's method; and (2) the method proposed by Pesaran *et al*. (2001). Table **2** reports the results.

Before interpreting the results of Table **2**, the following three notes are in order. First, in the Johansen procedure, after setting the cointegration rank to unity (r = 1), the diagnostic tests indicate that: (1) there is no autocorrelation in the residuals, since the p-values of the two *LM* statistics produced by the program are 0.31 and 0.81; (2) there exist strong *ARCH* effects (the p-values of the two *LM* statistics produced by the program are both 0.000); and (3) the normality assumption is also strongly rejected at the 1% level. Second, using this model, each variable was tested separately for weak exogeneity. The p-values of the relevant χ_1^2 statistic (produced by the program) are as follows: 0.04 for C_t, 0.01 for Y_t, and 0.12 for G_t. Based on this result, G_t was treated as weakly exogenous, and Equation (8) was re-estimated, thus obtaining the results of Table **2**. Note that imposing weak exogeneity of G_t lessens the *ARCH* effects (the p-values of the two *LM* statistics become 0.054 and 0.005), but the failure of normality remains a problem. According to Gonzalo (1994), however, normality is not crucial for the Johansen procedure. Third, in the case of the method of Pesaran *et al*. (2001), the dummy $D2001_t$ and the time trend turned out to be statistically insignificant, so they were dropped; the equation actually estimated is

$$\Delta C_t = \beta_0 + \beta_1 C_{t-1} + \beta_2 Y_{t-1} + \beta_3 G_{t-1} + \sum_{i=2}^{3} \varphi_i \Delta C_{t-i} + \psi_0 \Delta G_t + \sum_{i=2}^{3} \psi_i \Delta G_{t-i} + \omega_0 \Delta Y_t + \sum_{i=2}^{3} \omega_i \Delta Y_{t-i} + \varepsilon_t. \tag{9}$$

The parameters of interest are recovered as $b_1 = -\beta_2/\beta_1$ and $b_2 = -\beta_3/\beta_1$ (by setting all the first differences equal to zero and by normalizing with respect to C_{t-1}). Approximate standard errors of their estimators are obtained from the well-known formula for non-linear restrictions (Greene 2000:298-299). Note that in the context of Equation (9), testing the restrictions b_1 = 1 and b_2 = -1 amounts to testing the

Table 2: Estimation and Testing of Equations (8) and (9)

Equation, Method	$\hat{b}_1$	$\hat{b}_2$	$\hat{c}_1$	$\hat{c}_2$	H$_0$: $b_1 = 1$, $b_2 = -1$	Diagnostic tests
Equation (8), Johansen procedure	1.199*** (0.114)	-0.386* (0.265)	972.16* (573.95)	-33.99*** (12.47)	χ^2_2 = 10.40*** [0.006]	No autocorrelation: of order 1, χ^2_4 = 3.6 [0.458]; of order 2, χ^2_4 = 2.4 [0.671]; Normality: χ^2_4 = 33.9*** [0.000]; No ARCH: of order 1, χ^2_9 = 16.7* [0.054]; of order 2, χ^2_{18} = 37.0*** [0.005]
Equation (9), Pesaran *et al*. (2001)	0.963*** (0.061)	-0.439* (0.313)	–	–	χ^2_2 = 22.44*** [0.000]	No autocorrelation of order 1 to 4: χ^2_1 =1.8 [0.18], χ^2_2 = 1.9 [0.38], χ^2_3 =2.1 [0.56], χ^2_4 = 3.7 [0.45] Homoscedasticity: t = -0.1 [0.94] Normality: χ^2_2 = 6.9** [0.03]; Correct specification (RESET): second power of the fitted values, $F_{1, 245}$ = 3.0* [0.09]; second & third power, $F_{2, 244}$ = 1.8 [0.16]; second to fourth power, $F_{3, 243}$ = 1.6 [0.19]

Notes: (1) ***, **, * indicate statistical significance at the 1%, 5%, and 10% level; (2) standard errors are in parentheses underneath the estimated coefficients; *p*-values are in square brackets; (3) the values $\hat{b}_2$ = -0.386 (*t*-ratio = -1.47) and $\hat{b}_2$ = -0.439 (*t*-ratio = -1.40) are considered significant at the 10% level, because the model proposed in this paper points to one-sided alternatives; (4) in the case of the Johansen procedure, the diagnostic tests are those produced by *CATS in RATS*, v. 2.0; (5) in the case of the method of Pesaran *et al*. (2001), the tests for autocorrelation are standard Breusch-Godfrey tests (Greene 2000:541); the homoscedasticity test is a *t*-test on the slope coefficient in the regression of the squared residuals on the squared fitted values; and the normality test is the standard Bera-Jarque test.

restrictions $\beta_1 + \beta_2 = 0$ and $\beta_1 - \beta_3 = 0$, respectively. Note also that the "bounds test" for cointegration, reported in Part B of Table **1**, is a standard *F*-test of the hypothesis $\beta_1 = \beta_2 = \beta_3 = 0$, but with critical values obtained from Table CI(iii) Case III of Pesaran *et al*. (2001, p. 300). Thus, Equation (9) is a convenient workhorse for testing the hypotheses of interest. The tests are valid irrespective of whether the variables involved are I(0) or I(1), provided that there may be only one "levels relationship" involving the dependent variable (Pesaran *et al*. 2001).

The diagnostic tests reported in Table **2** and the notes made just before Equation (9) suggest that the reported coefficient estimates and the tests of the two restrictions can be considered reliable. According to these tests, both under the Johansen procedure and under the method of Pesaran *et al*. (2001), the joint hypothesis of interest can be rejected at the 1% level, because the restrictions $b_1 = 1$ and $b_2 = -1$ (tested jointly) are rejected. Testing these two restrictions individually reveals that, at the 5% level, the hypothesis $b_1 = 1$ cannot be rejected, since the *p*-values under the two methods are 0.08 and 0.55, whereas the hypothesis $b_2 = -1$ is rejected, since these two *p*-values are 0.02 and 0.04. Thus, *ex ante* full crowding out is rejected. Note also that the estimates of b_2 obtained from the two methods, -0.386 and -0.439, differ from zero only at the 10% level, and only because the alternative hypotheses are considered to be one-sided, in accordance with the model proposed in this paper.

4. SUMMARY AND CONCLUSIONS

The paper proposes a new approach to testing the hypothesis of *ex ante* full crowding out of government purchases by making only minimal assumptions and avoiding the complicated structures imposed by fully-fledged rationality. It only exploits the Ricardian assumption that a typical household incorporates the government's budget constraint into its own, and then tests the implications of this assumption as a joint hypothesis. In particular, the "combined" budget constraint implies an estimating equation that looks like a "consumption function," with income and government purchases as explanatory variables. Given that all of these variables are I(1), the joint hypothesis is that the "consumption function" is a cointegrating regression with coefficients of income and of government purchases 1 and -1, respectively.

This approach is applied to US quarterly data, 1947.1-2012.1. There is strong evidence for cointegration, but the above two coefficient restrictions are rejected. The real culprit for this rejection is the restriction $b_2 = -1$. In particular, the estimates of b_2 imply that an increase in government purchases by one dollar does not reduce private consumption by one dollar, as the model predicts, but only by about 40 cents. Thus, the hypothesis of *ex ante* full crowding out is rejected. As a policy conclusion, government purchases stimulate aggregate demand and can be used to fight a recession. The government should be careful in its spending, however. For if it finances its purchases by borrowing, its foreign debt is expected to rise, since the estimates of b_2 imply that private saving in the US will rise by less than the amount of the purchases, and so the government will have to turn to foreign borrowing.

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