Econometric Tests of Ricardian Equivalence: Results for Germany[1]

by

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Abstract

The approximate validity of the Ricardian equivalence proposition remains hotly disputed despite a large body of empirical work devoted to this issue. Contradictory conclusions arise at least partially from the fact that a large variety of very different tests has been applied to very different data sets. Moreover, instationarity properties of the data are not properly taken into account in some of the older literature, so that inference may be misguided. This paper collects the most important tests of Ricardian equivalence and applies them to a single data set of quarterly West German macro data. It turns out that there is hardly any evidence against the Ricardian proposition. Paradoxically, the theoretical model that implies Ricardian equivalence, is strongly rejected by the data.

JEL: E62, H31, C32

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1. Introduction

The theoretical case of Ricardian equivalence (REQ) is most easily demonstrated in a model of an infinitely lived representative agent. Under rather strong assumptions (e. g. fully rational, optimizing behavior, perfect capital markets and nondistortionary taxation), it is easy to show that *for a given level of government expenditures* the tax-debt composition does not have any real effects. The agent simply recognizes that debt always implies taxation in the future with the present value of future taxes being precisely equal to the amount of debt issued. Thus permanent income is unchanged and so are his optimal decisions.

Many plausible deviations from the baseline model are known to cause Ricardian equivalence to break down. Thus there seems to be a consensus in the literature that exact Ricardian equivalence cannot be expected to hold empirically. However, many people claim that REQ is a good and useful approximation to reality, while many others fiercely dispute this assertion. For instance, two rather comprehensive surveys of the literature by Bernheim (1987) and Seater (1993) derive the following polar conclusions:

"Taken together, the existing body of theory and evidence does not justify claims that government borrowing has little or no effect on the economy. Rather, I conclude that there is a significant likelihood that deficits have large effects on current consumption, and there is good reason to believe that this would drive up interest rates."

Bernheim (1987, p 264)

"Well designed tests based on consumption, economic growth, foreign trade, and exchange rates virtually unanimously suggest that Ricardian equivalence describes the data well; tests based on interest rates often do so. [...] Although some of the early empirical literature sent conflicting signals, recent work generally supports Ricardian equivalence."

Seater (1993, p. 182, p. 184)

A careful reading of the literature reveals that conflicting results on the validity of Ricardian equivalence are often attributable to problems of measurement and econometric methodology. While the former are sometimes difficult to handle and clear-cut solutions may not exist, progress in econometric theory has been rapid since the early 1980's, and much of the inference drawn in older time series studies is now known to be invalid. Moreover, although a wide battery of different tests exists, most researchers apply just one sort of test to a specific data set. As other studies often differ, not only by sample and data set, but also by the type of test employed, it is in generally difficult to assess how much evidence a given set of papers assembles in favor of or against Ricardian equivalence.

For empirical matters, Ricardian equivalence can sensibly be stated only with reference to a specific data set[2]. In this paper, I will use quarterly German macro data 1960.1-1994.4 (the territorial definition is West Germany throughout[3]) and subject them to the most important tests of REQ, partially updated to meet modern econometric standards. The definitions of the variables in the data set conform to their common usage in other applied German macro

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[2]Models always simplify reality, and it is hence of no surprise when a researcher finds a data set which is in some way in disaccord with a given model and its implications. The interesting question is the reverse one: For a given data set, researchers are invited to specify a model and test whether the model can be rejected with this particular data set.

[3]See the appendix to this paper for a brief description of the data.

studies, since it is for many of these papers (which study unrelated questions of public policy) that REQ might be a useful feature to assume.

The theoretical setup of Ricardian equivalence is well known to most economists, numerous excellent expositions exist. (In addition to the above surveys, see Leiderman and Blejer (1988)). I will therefore not bore the reader with yet another sketch of the theory. Rather, I will take estimable equations from the literature and move directly into the empirical investigation, which is truely at the heart of the debate. The paper is therefore organized as follows: Section 2 reports results on both the standard and the consolidated consumption function tests of Ricardian equivalence. Section 3 also studies consumers' behavior, but in an Euler equation setting. This part of the paper allows for finite horizons and liquidity constraints. Section 4 presents tests on interest rates and Section 5 tests on exchange rates. An overall evaluation of the evidence produced is given in the Conclusions.

2. Consumption Function Tests

Tests based on the specification of "structural" consumption functions are certainly the most popular devices in testing Ricardian equivalence, although some researchers (Barro (1989)) voice serious doubts about the quality of this approach. Some of the most important papers in this line of research are due to Feldstein (1982), Kormendi (1983), Aschauer (1985), Seater and Mariano (1985), Blinder and Deaton (1985), Evans (1988), and Haug (1990). These studies typically specify a consumption function similar to

$$C_t = \beta_0 + \beta_1 Y_t + \beta_2 W_t + \beta_3 D_t + \beta_4 G_t + \beta_5 T_t + \beta_6 SS_t + \beta_7 TR_t + error, \qquad (1)$$

where C_t is a measure of consumption, Y_t is current income, W_t is household wealth, D_t is government debt, G_t is a measure of government purchases, T_t are total tax proceeds, SS_t is a measure of social security claims, and TR_t are government transfers to private households.

Before turning to the precise definitions of these variables, it is worth mentioning that there are (at least) two econometric complications associated with : First, there is a simultaneity problem as the regressors are correlated with the error term. Hence instrumental variables (IV) estimates are required to obtain consistent estimates. Second, some of the variables are likely to be integrated of order one, I(1), so that regression results make sense only if the dependent variable is cointegrated with the regressors. Even in the case of cointegration, the (asymptotic) distribution of the regression coefficients is nonstandard. Essentially, estimation of (1) is the first step of Engle and Granger's (1987) test for cointegration; this step, while yielding superconsistent estimates, is known to result in rather unreliable estimates in finite samples, since the specification does not take the short run dynamics of the system into account (see, e. g., Inder (1993)).

Feldstein (1982) is to be credited for introducing IV-estimates in this strand of literature, and Kormendi (1983) is an early example for sensing the jeopardies of regressions with instationary variables. Kormendi chose to estimate his (consolidated) consumption function in first differences rather than levels. While this is not necessarily recommended from today's econometric knowledge (first differencing may actually induce misspecification since the long run equilibrium relationship is neglected), I will nevertheless follow his footprints and work with first differenced data first. This decision is partly motivated by the non-availability of some stock variables and partly by the fact that systems estimates below indicate that consumption might well be weakly exogenous, in which case the differenced form of could be correctly specified even in the case of cointegration[4]. I will elaborate on these points in the following sections.

Let us first turn to the data: I use quarterly, seasonally non-adjusted data 1960.1-1994.4 from the system of national accounts maintained by the Deutsches Institut für Wirtschaftsforschung (DIW), Berlin. (This data set also contains the data on interest rates, exchange rates and terms of trade, that will be used in Sections 4 and 5). I complement these data by time series for monetary wealth supplied by the Deutsche Bundesbank (1992) and extrapolated to (seasonally non-adjusted) quarterly data 1960.1-1994.4 by myself[5]. The territorial definition of all data pertains to West Germany prior to unification and the price base for all real variables is 1991.

In the spirit of the traditional consumption function tests (as opposed to Kormendi's consolidated approach) I use real private consumption for C_t. Y_t is represented by real GDP and G_t is real government consumption plus real government net investment[6]. T_t denote real domestic tax revenues and TR_t real domestic transfers. Since Germany's social security systems are overwhelmingly on a pay-as-you-go basis, SS_t is given by real domestic contributions to the social security systems, rather than a measure of the stock of social security wealth. The change in household wealth ΔW_t is by definition equal to the change in private enterprise wealth minus the change in government wealth minus the change in wealth of the rest of the world (ROW). Since the change in government wealth is equal to minus the change in government debt D_t, I split ΔW_t into $\Delta V_t + \Delta D_t$, where ΔV_t is equal to the change in private enterprise wealth minus the change in ROW-wealth.

Note that under the null hypothesis of Ricardian equivalence D_t is not net wealth and V_t therefore denotes all household wealth. One would thus expect $\beta_2>0$ and $\beta_3=0$. Further, increasing G_t would imply higher taxes and lower permanent income, thus it is often stated that we should have $\beta_4<0$. Since G_t fully captures the level of government absorption of resources, the present tax load, social security contributions and transfers do not affect the the optimal consumption decisions of a Ricardian consumer, thus the null implies $\beta_5=\beta_6=\beta_7=0$[7].

By contrast, in a traditional Keynesian setting, government bonds would be perceived as net wealth, and government spending should increase private consumption, thus we would have $\beta_3>0$ and $\beta_4>0$. A higher percentage of tax financing cet. par. lowers consumption, as do increased social security contributions, whereas transfers tend to support consumption. Hence the alternative hypothesis would imply $\beta_5<0$, $\beta_6<0$, and $\beta_7>0$. Clearly, under both hypotheses β_1 and β_2 should be positive.

I estimate equation (1) in differenced form, using three seasonal dummies S1, S2, and S3 and a reunification dummy D903 which is zero prior to German currency union and one thereafter[8]. I also allow for an MA(1) term to capture errors in temporal aggregation and an

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[4]This reasoning basically treats variables as weakly exogenous, when their first differences do not depend on error correction terms in a vector error correction model. Strictly speaking, the conditions for weak exogeneity are somewhat weaker, cf. Banerjee, Dolado, Galbraith, and Hendry (1993).

[5]In inter- and extrapolating the Bundesbank data, I basically used the data on change in monetary wealth contained in the DIW's quarterly national accounts. All my data are available upon request.

[6]If gross rather than net government investment were used, regression results for private consumption could not be expected to be invariant with respect to public capital stock consumption.

[7]Unless these measures have distributional effects. See Kormendi (1983) and Haug (1990).

[8]I do allow for a reunification dummy despite of the fact that I use data confined to the territorial definition of the former West Germany, since even from a purely West German perspective the reunification doubtlessly

MA(4) term to capture remaining stochastic seasonality[9]. The estimation methods is two-stage-least squares (2SLS), with the deterministic variables and the stochastic regressors lagged for four quarters serving as instruments[10]. The results are given in Table 1:

Table 1

2SLS Estimates of Structural Consumption Function: Dependent Variable is ΔC_t

Variable	Coefficient	Std. Error	t-Statistic	Prob.
CNST	12.40	4.028	3.078	0.003
ΔY_t	0.226	0.091	2.488	0.014
ΔV_t	-0.024	0.081	-0.295	0.768
ΔD_t	-0.011	0.035	-0.318	0.751
ΔG_t	0.317	0.475	0.667	0.506
ΔT_t	0.008	0.270	0.029	0.977
ΔSS_t	0.323	0.217	1.486	0.140
ΔTR_t	0.022	0.341	0.065	0.948
S1	-25.24	7.61	-3.32	0.001
S2	-7.10	2.47	-2.88	0.005
S3	-12.49	3.19	-3.91	0.000
ΔD903	1.37	2.91	0.469	0.640
MA(1)	-0.317	0.092	-3.427	0.001
MA(4)	0.406	0.096	4.242	0.000
R^2	0.981		Mean dep. var	2.049407
$\overline{R}^2$	0.979		S.D. dep. var.	18.80
S.E. of regr.	2.716		Akaike	2.10
SSR	892.52		Schwarz	2.40
F-statistic	478.27		Durbin-Wats.	2.06
Wald-Tests on Redundant Variables: $\beta_3=\beta_5=\beta_6=\beta_7=0$				
	F=1.289 P(F)=0.278		χ^2=5.158 P(χ^2)=0.271	

Here, as in all regressions in this paper, White's (1980) heteroskedasticity consistent standard errors have been used. It turns out that the only significant "structural" variable is the change

represented a large macroeconomic shock that likely affected the time series behavior of the macroeconomic aggregates.

[9]I did not explicitly test for seasonal unit roots, but if there were any, then we should see a root close to unity in the MA(4) polynomial. Since no such evidence obtains, we may either suppose the absence of seasonal unit roots or seasonal cointegration between dependent and independent variables.

[10]As the estimation method is 2SLS, linear hypothesis tests have been computed in the modification suggested by Startz (1983).

in income. In particular, for those variables with a coefficient of zero under REQ, Wald tests of the joint hypothesis cannot be rejected. Looking at the t-statistics, the insignificance of the wealth term ΔV_t is not in accord with either hypothesis, even though this is probably not too troubling, since wealth effects are usually suspected to be rather small. Insignificant changes in public debt, taxes, social security contributions, and transfers are in line with Ricardian equivalence, but incompatible with the Keynesian alternative.

Finally, the insignificance of the change in net government absorption seems incompatible with either hypothesis. On second thought, however, the insignificance is hard to interpret under the null of Ricardian equivalence, since a careful analysis shows that $\beta_4 < 0$ may actually fail to hold under this hypothesis. The reason is a misspecification of equation (1) in the sense that under REQ consumption should depend on permanent income, not current income. Since Baxter and King's (1993) analysis of the neoclassical growth model showed that an increase in government absorption may cause a more than one-to-one increase in permanent output and thus an increase in consumption, neglecting permanent income in equation (1) will induce a positive correlation between private consumption and government absorption due to an omitted variables problem. This positive correlation counteracts the negative correlation between the two variables expected under REQ. Hence, the insignificance of the change in government absorption might well be accommodated with REQ. Note that a similar argument could be made for the alternative hypothesis. Yet here the positive correlation would be reinforced by the omitted variables problem, hence the insignificant coefficient is not compatible with the alternative.

In an important modification of the above specification Kormendi (1983) argued in favor of a so-called "consolidated approach" to consumption, which basically treats private and government consumption as jointly determined consumption expenditures of a representative agent. Taking his point of view into account and simultaneously checking the robustness of the results displayed in Table 1 I reran the above regression with the modification that the dependent variable ΔC_t is replaced by $\Delta C_t + \Delta GC_t$, where GC_t is government consumption, and ΔG_t is replaced by ΔGI_t, where GI_t is net government investment. The results, given in Table 2, are very similar to those of the standard approach[11].

[11]I have deleted the dummy regressor ΔD903, since it was not significant. Including this regressor as in Table 1 leads to virtually the same results, except that the P-value of ΔY_t is 0.07.

Table 2

2SLS Estimates of Structural Consumption Function: Dependent Variable is $\Delta C_t + \Delta GC_t$

Variable	Coefficient	Std. Error	t-Statistic	Prob.
CNST	15.85	3.25	4.88	0.000
ΔY_t	0.330	0.128	2.580	0.011
ΔV_t	0.034	0.119	0.287	0.775
ΔD_t	0.075	0.056	1.341	0.182
ΔGI_t	0.888	1.092	0.813	0.418
ΔT_t	0.324	0.300	1.082	0.281
ΔSS_t	0.140	0.322	0.435	0.665
ΔTR_t	0.333	0.399	0.834	0.406
S1	-29.73	8.96	-3.32	0.001
S2	-14.43	8.22	-1.76	0.082
S3	-17.87	2.46	-7.27	0.000
MA(1)	-0.242	0.080	-3.035	0.003
MA(4)	0.413	0.094	4.370	0.000
R^2	0.977		Mean dep. Var	2.668
$\overline{R}^2$	0.975		S.D. dep. var.	24.83
S.E. of regr.	3.956		Akaike	2.84
SSR	1909.5		Schwarz	3.12
F-statistic	428.6		Durbin-Wats.	2.07
Wald-Tests on Redundant Variables: $\beta_3=\beta_5=\beta_6=\beta_7=0$				
	F=1.607 P(F)=0.177		χ^2=6.428 P(χ^2)=0.169	

It is, of course, of highest interest, to include level variables in this sort of analysis. However, integrating ΔV_t and ΔD_t does not result in reasonable stock variables, since the integrated series do not take changes in the market value of the assets into account. As time goes by, the resulting error will be substantial. In the subsequent analysis, I will therefore use market-value adjusted data of the stock variables provided by the Deutsche Bundesbank[12]. Due to market-value adjustment, the first differences of these stock variables do not coincide with the corresponding flow data of the system of national accounts which I have used above. This is somewhat unfortunate, since the analysis of cointegrated systems requires both the use of level variables and of their first differences. In order to maintain a consistent data handling in the cointegration analysis, I will use the first differences of the (market-value adjusted) stock variables in place of the flows ΔV_t and ΔD_t from now on. To make the distinction explicit, variables referring to market value adjusted series are henceforth characterised by a tilde, i. e. $\tilde{V}_t$, $\tilde{D}_t$, $\Delta\tilde{V}_t$, and $\Delta\tilde{D}_t$.

[12]So-called "Geldvermögensrechnung".

Let us first have a look at the cointegration properties of the variables. (All cointegration tests that follow include seasonal dummies and D903 as exogenous series). Looking at the full system first, i. e. at the eight level variables C_t, Y_t, $\tilde{V}_t$, G_t, $\tilde{D}_t$, T_t, SS_t, and TR_t, we find evidence of two, possibly even three cointegrating vectors, cf. Table 3[13]. This result, however, is hardly surprising, since it is quite common to find cointegration in high dimensional systems. Also, the mere existence of cointegration is hard to interpret in terms of REQ, and the estimated cointegrating vectors are probably not very helpful, as the appropriate normalisation is not clear and the coefficient estimates may be rather unprecise due to the large number of variables.

Table 3

Johansen Cointegration Tests
(Intercept in Cointegrating Equation and Test VAR)

Endogenous Variables: C_t, Y_t, $\tilde{V}_t$, G_t, $\tilde{D}_t$, T_t, SS_t, and TR_t
Exogenous Variables: *CNST, S1, S2, S3, D903*; Lags 1, 2, 4

Eigenvalue	LR (Trace-Stat.)	5 Percent Critical Value	1 Percent Critical Value	Hypothesized No. of CE(s)
0.507	247.10 **	156.00	168.36	None
0.351	151.60 **	124.24	133.57	At most 1
0.209	93.31	94.15	103.18	At most 2
0.192	61.64	68.52	76.07	At most 3
0.107	32.78	47.21	54.46	At most 4
0.070	17.50	29.68	35.65	At most 5
0.037	7.67	15.41	20.04	At most 6
0.018	2.52	3.76	6.65	At most 7
*=5% significant, **=1% significant				

It is therefore more interesting to consider two four-dimensional subsystems, the first of which consists of the variables that should be cointegrated under REQ, i. e. C_t, Y_t, $\tilde{V}_t$ and G_t, and the second comprising the remaining variables $\tilde{D}_t$, T_t, SS_t, and TR_t, which, under REQ, should not have any (long run) impact on private consumption. If the total number of cointegrating vectors in the two subsystems equals the number of cointegrating vectors in the full system, then the two subsystems can be thought of as being characterised by independent long-run dynamics, i. e. independent stochastic trends, and this finding would clearly be supportive of REQ.

As Tables 4 and 5 show, there is actually evidence for a total of three (possibly even four) cointegrating vectors in the two subsystems. In view of the tests for the full system, it is hence likely that there is complete separation between the two subsystems and therefore evidence for

[13]Note that there is no identification problem. The Johansen procedure identifies the basis of the cointegration space and all I have to show under both REQ and the Keynesian alternative is that there exists a basis of this space which can be meaningful interpreted with respect to one of these hypotheses.

Ricardian equivalence in the long run. This conclusion is supported by the estimated coefficients of the cointegrating vector for the first subsystem. In the long-run, consumption seems to increase approximately one-to-one with income, but to decrease approximately one-to-one with government absorption. This behavior is precisely the prediction of REQ! However, what is not in line with the theory is the long run coeffcient of the wealth variable $\tilde{V}_t$. This coefficient is significant and has the wrong sign, thus consumption is counterintuitively and counterfactually predicted to fall with rising wealth. This result is somewhat troubling, although the coefficient estimate is rather small and might possibly just account for the fact that the coefficient estimate for income exceeds unity by approximately the same amount.

Table 4

Johansen Cointegration Tests

(Intercept in Cointegrating Equation and Test VAR)

Endogenous Variables: C_t, Y_t, $\tilde{V}_t$, and G_t

Exogenous Variables: *CNST, S1, S2, S3, D903*; Lags 1-4

Eigenvalue	LR (Trace-Stat.)	5 Percent Critical Value	1 Percent Critical Value	Hypothesized No. of CE(s)
0.320	80.76 **	47.21	54.46	None
0.117	28.59	29.68	35.65	At most 1
0.077	11.72	15.41	20.04	At most 2
0.007	0.91	3.76	6.65	At most 3
*=5% significant, **=1% significant				
Estimated coefficients of cointegrating vector (standard errors)				
C_t	Y_t	$\tilde{V}_t$	G_t	*CNST*
1	-1.087 (0.123)	0.071 (0.017)	0.92 (0.33)	51.70

Looking at the estimated cointegrating vectors for the second subsystem we see that these also make sense. In particular, the second cointegrating vector basically states that taxes plus social security contributions approximately cover all transfer payments, hence government absorption is largely financed by issuing debt. The first vector reflects the fact that social security contributions alone are not sufficient to finance transfers and that the share of transfers not covered by social security contributions grows at roughly the same rate as total government debt.

Table 5

Johansen Cointegration Tests
(Intercept in Cointegrating Equation and Test VAR)

Endogenous Variables: $\tilde{D}_t$, T_t, SS_t, and TR_t
Exogenous Variables: *CNST*, *S1*, *S2*, *S3*, *D903*; Lags 1 ,2 ,4

Eigenvalue	LR (Trace-Stat.)	5 Percent Critical Value	1 Percent Critical Value	Hypothesized No. of CE(s)
0.291	83.25 **	47.21	54.46	None
0.163	36.81 **	29.68	35.65	At most 1
0.088	12.79	15.41	20.04	At most 2
0.003	0.37	3.76	6.65	At most 3
*=5% significant, **=1% significant				
Estimated coefficients of cointegrating vector (standard errors)				
$\tilde{D}_t$	T_t	SS_t	TR_t	*CNST*
1	0	73.04 (25.35)	-66.63 (26.54)	371.88
0	1	1.26 (0.82)	-2.37 (0.86)	-9.34

The long-run time series properties of the data are hence quite well in line with Ricardian equivalence. I will thus now turn my attention to the short-run dynamics. For that purpose, let us denote the cointegrating combination defined by the cointegrating vector in Table 4 as EC0, and the cointegrating combinations defined by the first and the second cointegrating vector in Table 5 by EC1 and EC2, respectively. I now regress the change in consumption on these error correction terms lagged one quarter and on the simultaneous changes in the other "structural" variables. The estimation method is again 2SLS with contemporaneous variables being instrumented by their own past one year ago.

The second column of Table 6 gives the results of these estimates under the null of REQ, i. e. by including just the regressors with nonzero coefficients under Ricardian equivalence. As before, changes in real wealth and government absorption turn out to be insignificant, while the change in current income is an important determinant of the change in consumption. The error correction term EC0(-1) is also insignificant.

Columns 3 to 7 test various alternative specifications in each of which a single additional regressor is added to the basic regression of column 2. According to the estimation results under these alternative specifications, the error correction terms of either subsystem are never significant, suggesting that the change in consumption is actually weakly exogenous. Thus, despite cointegration between the level variables, a specification in first differences only is appropriate, lending support to my analysis above. Further, the change in income is almost always significant, with one exception where the P-value is 0.06. The real wealth variable is never significant, and the change in government absorption is significant just once with a positive coefficient. While this one significant statistic is probably an outlier, it is indeed compatible with REQ, while the insignificant statistics for government absorption are not compatible with the Keynesian alternative, as pointed out above. The additional regressors $\Delta\tilde{D}_t$, ΔT_t, ΔSS_t, and ΔTR_t, finally, are insignificant throughout, thus not lending support to the Keynesian alternative either.

Similar results are obtained, when the basic regression is specified even more parsimoniously. Taking as the point of departure just those regressors that are significant in the specification of Column 2 in Table 6 and testing alternative specifications with just one additional regressor, yields regression results as displayed in Table 7. Income and the MA-terms are significant at the 1% level throughout. All other structural regressors are insignificant, with the exception of the change in social security contributions, which has a P-value of 4%. As the associated coefficient is positive, this estimation result, if taken seriously, is not in accord with either REQ or the Keynesian alternative. It might hint at weak distributional effects or it might be just a statistical artefact. In fact, if all the t-tests in Tables 6 and 7 were independent, we should expect that one out of twenty statistics is significant at the 5% level even when all coefficients are truely zero.

Table 6

Various Error Correction Specifications for the Change in Consumption, Part I

	Basic Regression	Modifications of Basic Regression					
EC0(-1)	-0.022 (0.034)	-0.032 (0.037)	-0.050 (0.045)	-0.026 (0.035)	-0.030 (0.039)	-0.015 (0.036)	-0.027 (0.033)
EC1(-1)	-	-0.001 (0.001)	-	-	-	-	-
EC2(-1)	-	-	-0.037 (0.031)	-	-	-	-
ΔY_t	0.230 * (0.091)	0.223 * (0.094)	0.258 ** (0.050)	0.232 * (0.092)	0.189 (0.103)	0.237 ** (0.074)	0.214 * (0.100)
$\Delta \tilde{V}_t$	-0.037 (0.023)	-0.041 (0.026)	-0.006 (0.011)	-0.034 (0.023)	-0.058 (0.035)	0.004 (0.034)	-0.041 (0.027)
$\Delta \tilde{D}_t$	-	-	-	-0.016 (0.029)	-	-	-
ΔG_t	0.350 (0.401)	0.373 (0.417)	0.325 * (0.150)	0.406 (0.441)	0.439 (0.425)	0.285 (0.410)	0.371 (0.424)
ΔT_t	-	-	-	-	0.008 (0.099)	-	-
ΔSS_t	-	-	-	-	-	0.298 (0.224)	-
ΔTR_t	-	-	-	-	-	-	0.170 (0.187)
MA(1)	-0.345 ** (0.091)	-0.329 ** (0.091)	-0.307 ** (0.103)	-0.347 ** (0.091)	-0.307 * (0.101)	-0.328 ** (0.094)	-0.364 ** (0.091)
MA(4)	0.353 ** (0.092)	0.338 ** (0.097)	0.393 ** (0.010)	0.381 ** (0.092)	0.313 * (0.105)	0.382 ** (0.096)	0.389 ** (0.087)
Q(16)	15.42	15.19	20.45	16.06	12.68	18.98	14.83
*=5% significant, **=1% significant. Standard errors in parentheses. Estimation results for constant and dummies are suppressed.							

Table 7

Various Error Correction Specifications for the Change in Consumption, Part II

	Basic Regression	Modifications of Basic Regression								
EC0(-1)	-	-0.027 (0.035)	-	-	-	-	-	-	-	-
EC1(-1)	-	-	-0.000 (0.001)	-	-	-	-	-	-	-
EC2(-1)	-	-	-	-0.008 (0.025)	-	-	-	-	-	-
ΔY_t	0.288 ** (0.075)	0.298 ** (0.077)	0.288 ** (0.076)	0.286 ** (0.078)	0.288 ** (0.078)	0.288 ** (0.081)	0.257 ** (0.090)	0.280 ** (0.081)	0.269 ** (0.082)	0.284 ** (0.079)
$\Delta \tilde{V}_t$	-	-	-	-	-0.006 (0.009)	-	-	-	-	-
$\Delta \tilde{D}_t$	-	-	-	-	-	0.015 (0.021)	-	-	-	-
ΔG_t	-	-	-	-	-	-	0.318 (0.214)	-	-	-
ΔT_t	-	-	-	-	-	-	-	0.066 (0.064)	-	-
ΔSS_t	-	-	-	-	-	-	-	-	0.303 * (0.143)	-
ΔTR_t	-	-	-	-	-	-	-	-	-	0.151 (0.160)
MA(1)	-0.320 ** (0.08)	-0.301 ** (0.101)	-0.319 ** (0.083)	-0.317 ** (0.084)	-0.324 ** (0.091)	-0.319 ** (0.083)	-0.331 ** (0.083)	-0.297 ** (0.089)	-0.310 ** (0.085)	-0.336 ** (0.083)
MA(4)	0.426 ** (0.09)	0.426 ** (0.095)	0.425 ** (0.095)	0.429 ** (0.097)	0.414 ** (0.092)	0.411 ** (0.092)	0.393 ** (0.089)	0.420 ** (0.094)	0.395 ** (0.094)	0.438 ** (0.089)
Q(16)	22.30	21.02	19.67	18.40	21.06	22.22	21.64	23.60	21.17	20.93

*=5% significant, **=1% significant. Standard errors in parentheses. Estimation results for constant and dummies are suppressed.

Thus, overall, the results I have presented so far are quite supportive of Ricardian equivalence. An analogous analysis can be done for the consolidated approach. The results turn out to be quite similar and are suppressed here to save space.

3. Euler Equation Tests

Testing Ricardian equivalence by estimating "structural" consumption functions has been subject to the critique that these functions, for instance in their use of current rather than permanent income, are incompatible with rational expectations and utility maximizing consumers. These features, however, are an indispensable cornerstone of REQ. Thus Flavin (1987), for instance, argues that the first order conditions from consumer maximization are a necessary condition for REQ to hold, and tests of Ricardian equivalence should therefore be based on Euler equation approaches.

Such tests have in fact been developed, and often have been nested in Blanchard's (1985) model of finite horizons[14]. In that model, a fraction μ of the population dies each period. Due to finite lifetimes, an agent may not live up to the point where the government pays back its debt, and since people who die are assumed not to care about the welfare of those that survive, Ricardian equivalence breaks down unless $\mu = 0$, i. e. unless all agents live infinitely long.

Blanchard's model implies a consumption function of the form

$$C_t = \alpha\left[(1+r)A_{t-1} + \sum_{j=0}^{\infty}\left(\frac{1-\mu}{1+r}\right)^j E_t Y_{t+j}^d\right], \tag{3}$$

where α is the propensity to consume out of total wealth, r is the constant real rate of return on assets, A_{t-1} is the stock of real assets outstanding at the end of period *t-1*, and Y_t^d is real after tax labor income. Using the aggregate budget constraint

$$A_t = (1+r)A_{t-1} + Y_t^d - C_t, \tag{4}$$

one can write consumption as a function of its own past and of nonhuman wealth,

$$C_t = \left(\frac{1+r}{1-\mu}\right)(1-\alpha)C_{t-1} - \alpha\mu\frac{1+r}{1-\mu}A_{t-1} + \alpha\varepsilon_t, \tag{5}$$

where ε_τ is an expectational error. This solution was used by Evans (1988). Haque (1988) derives an equivalent solution which depends only on consumption and disposable labor income:

$$C_t = (1+r)\left(1-\alpha+\frac{1}{1-\mu}\right)C_{t-1} - \frac{(1+r)^2}{1-\mu}(1-\alpha)C_{t-2} - \alpha\mu\frac{1+r}{1-\mu}Y_{t-1}^d + \alpha\varepsilon_t - \alpha(1+r)\varepsilon_{t-1} \tag{6}$$

A third equivalent solution is due to Hayashi (1982), taking the form

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[14] See Cardia (1997) for a recent simulation study on the merits of these tests.

$$C_t = \frac{(1+r)}{(1-\mu)}\left(1-\alpha(1-\mu)\right)C_{t-1} - \alpha\mu\frac{(1+r)^2}{1-\mu}A_{t-2} - \alpha\mu\frac{1+r}{1-\mu}Y_{t-1}^d + \alpha\varepsilon_t. \tag{7}$$

All three solutions have the property that under infinite horizons, i. e. REQ, μ=0 implies that current consumption is a function of lagged consumption only. Under the alternative of finite horizons (where REQ breaks down), the additional regressors should be significant.

I will start the analysis with Johansen's test for cointegration. Using real disposable income minus real after tax income from entrepreneurial activities as the measure for Y_t^d, real private consumption expenditures for C_t, and real household wealth for A_t, the trace-statistic for a VAR(4) with seasonal dummies is 28.98, while the 5% critical value is 29.68. While the null of no cointegration cannot be rejected with this statistic, the test statistic is close enough to the critical value to justify an analysis of equations (5) to (7) in levels, i. e. assuming cointegration. In fact, if the null of no cointegration were true, then the whole model of a rational optimizing consumer would be rejected by the data, as the model explicitly implies cointegration.

Estimation results are given in Table 8, where I have used OLS for equations (5) to (7), since all regressors are lagged for at least one period so that there is no simultaneity bias between regressors and the contemporaneous innovation[15]. I have again included an MA(1) term to capture temporal aggregation effects as well as seasonal dummies and an MA(4) term to account for seasonality[16]. Coefficient estimates for these regressors are not given to save space, but they are typically highly significant and the roots of the MA-polynomial are well

[15]However, since I specify an MA-disturbance, there may be a simultaneity problem in finite samples since the regressors could be correlated with the innovations lagged one and four periods. Using instrumental variables may then require to lag the regressors by as much as five periods. As these are likely to be rather poor instruments (and better ones not readily available), the analysis of Nelson and Startz (1990) suggests that OLS estimation may in fact be preferable to IV.

[16]Thus I estimate all Euler equations with MA-specifications even though only in equation (6) an MA structure is required by the model. This seems appropriate since there is no a priori reason to suppose that the planning period of the representative individual coincides with a quarter and since the theoretical model does not allow for seasonality. The case of white noise residuals is, of course, nested in the MA-specification.

inside the invertibility region. In most cases the MA-terms are significant; an important exception is noted below. Significance levels are computed from a standard t-distribution[17].

[17]This is, probably not even asymptotically, the correct distribution for the (instationary) regressors. While I am not aware of any work deriving the appropriate distribution of the t-statistics for the above regressions (and similar ones that are to follow below), it is likely that the true critical values are similar to the critical values published by Banerjee, Dolado and Mestre (1994). If this were so, then no regressor except lagged consumption would be significant at the 5%-level and hence the regressions (5) to (10) would not be able to produce evidence against REQ. However, the framework of Banerjee, Dolado and Mestre is not precisely the same as the model used here and so the applicability of their critical values is less than clear.

Table 8

Euler Equation Estimates

	Coefficient estimates (standard errors in parentheses)							Q-statistic (P-value)	Q-statistic (P-value)
Eq.	C_{t-1}	C_{t-2}	A_{t-1}	A_{t-2}	Y_t^d	Y_{t-1}^d	Y_{t-2}^d	Q(8)	Q(16)
(5)	0.930 ** (0.031)	-	0.011 ** (0.004)	-	-	-	-	88.850** (0.000)	140.03** (0.000)
(6)	0.844 ** (0.154)	0.160 (0.152)	-	-	-	-0.006 (0.026)	-	16.430* (0.037)	20.49 (0.115)
(7)	0.935 ** (0.033)	-	-	-0.011 (0.015)	-	0.011 * (0.005)	-	95.862** (0.000)	153.34** (0.000)
(8)	0.885 ** (0.057)	-	0.007 * (0.003)	-	0.086 * (0.033)	-0.042 (0.031)	-	7.703 (0.261)	9.35 (0.808)
(9)	0.838 ** (0.159)	0.171 (0.162)	-	-	0.115 ** (0.037)	-0.023 (0.046)	-0.102 ** (0.037)	3.432 (0.753)	7.45 (0.916)
(10)	0.880 ** (0.055)	-	-	-0.008 * (0.003)	0.088 ** (0.033)	-0.043 (0.031)	-	7.049 (0.316)	8.76 (0.846)

*=5% significant, **=1% significant

In principle, the estimates from equations (5) to (7) should yield the same conclusions. They in fact do, but in a rather surprising way. While the coefficient estimates from equations (5) and (7) seem to suggest evidence unfavorable to REQ, these equations suffer from severe autocorrelation in the residuals. As the Euler-equation error should be white noise under rational expectations, the presence of strong autocorrelation simply indicates that neither model of consumer optimization (finite or infinite horizons) actually fits the data.

Autocorrelation is much less of a problem in equation (6), where out of the first 36 lags, lag 8 is the only lag at which the Q-statistic has a P-value lower than 5%. The coefficient estimates, however, are not in line with the theoretical model. For instance, C_{t-2} is not significant, although it should display a negative coefficient. Further, the MA(1) term is not significant, although it is explicitly called for by the model. Finally, the coefficient of C_{t-1} should be larger than one if the model were true, but the converse is the case. Thus, all three equations do indeed yield the same conclusion: The simple model of rational consumer optimization is not supported by the data[18].

Himarios (1995) argues that an appropriate modification of the model might be achieved by allowing for liquidity constraints as in Campbell and Mankiw (1990)and he presents empirical evidence (US data) in favor of this hypothesis. Relaxing the assumption of perfect capital markets (which is indeed crucial for REQ), and assuming that a fraction λ of the population is liquidity constrained, equations (5) to (7) become:

$$C_t = \frac{1+r}{1-\mu}(1-\alpha)C_{t-1} - \alpha\mu\frac{1+r}{1-\mu}A_{t-1} + \lambda Y_t^d - \lambda\frac{1+r}{1-\mu}(1-\alpha)Y_{t-1}^d + u_t \tag{8}$$

$$\begin{aligned} C_t = {} & \left(\frac{1+r}{1-\mu}\right)\left(1+(1-\alpha)(1-\mu)\right)C_{t-1} - \frac{(1+r)^2}{1-\mu}(1-\alpha)C_{t-2} + \lambda Y_t^d \\ & - \left(\frac{1+r}{1-\mu}\right)\left\{\alpha\mu + \lambda[2-\alpha-\mu]\right\}Y_{t-1}^d + \lambda(1-\alpha)\frac{(1+r)^2}{1-\mu}Y_{t-2}^d + \eta_t \end{aligned} \tag{9}$$

$$C_t = \frac{(1+r)}{(1-\mu)}\left(1-\alpha(1-\mu)\right)C_{t-1} - \alpha\mu\frac{(1+r)^2}{1-\mu}A_{t-2} + \lambda Y_t^d - \frac{1+r}{1-\mu}\left[\lambda - \alpha(\lambda-\mu)\right]Y_{t-1}^d + u_t. \tag{10}$$

Here, u_t is white noise and $\eta_t = u_t - (1+r)u_{t-1}$. I estimate equations (8) to (10) by 2SLS, where contemporaneous after tax labor income is instrumented by its own value lagged four quarters. The results are also given in Table 8.

Note first that autocorrelation in the residuals has vanished with the introduction of liquidity constraints. The estimated coefficients, however, cannot be reconciled with either the null of infinite horizons or its alternative. In equation (8), for instance, the coefficient of household wealth is significantly positive, while it should be zero under the null, and negative under the alternative. In addition, lagged income is insignificant as it should be under the null, while current income is significantly positive as it should be with imperfect capital markets. Hence, with conventional distribution theory in mind, neither the null nor the alternative is supported by the estimation results for equation (8).

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18 Note that a multicountry Euler-equation study by Evans (1993) also produces some evidence against REQ for West Germany. However, Evans only reports the t-statistics of the coefficients of interest; he does not give standard specification test statistics. It is hence unclear from his publication if the alternative of the finite lifetime Blanchard model is in fact compatible with the data.

Results for equation (9) do not fit the model either, since for instance the coefficient for C_{t-1} should be larger than one, while the point estimate is substantially smaller. However, due to a large standard error, one could not discard the possibility that the true coefficient is in fact larger than one. But it is also troubling that the coefficient of C_{t-2} is insignificant (in contradiction to both the null and the alternative). And, more importantly, the coefficient of Y^d_{t-2} is significantly negative, which is also not in accord with either hypothesis. In equation (10), finally, we see that the coefficient for lagged income is not significantly different from zero. But a zero coefficient for lagged income implies $\lambda=-\alpha\mu/(1-\alpha)$, i. e. λ must be negative, which is in contradiction to both the significantly positive coefficient estimate for contemporaneous income and the model interpretation of λ as the percentage of liquidity constrained households.

Taken together, it seems that the Euler equation tests do not provide evidence against the Ricardian null hypothesis by favoring a suitably modelled alternative. Rather, they suggest that there is something wrong with the model of an optimizing, rational representative consumer in the first place. So thus far we are left with the rather paradoxical result that the theoretical model which is usually employed to motivate the Ricardian hypothesis is soundly rejected by the data, while the approximate validity of Ricardian equivalence as an empirical phenomenon can be upheld in analyses with consumption data. This finding is similar to an analysis with Canadian data by Haug (1996), who finds no evidence against REQ in Blanchard's framework, but strongly rejects Blanchard's perpetual youth model in the first place. Moreover, Cardia (1997) shows that tests of REQ in Blanchard's model may lack robustness properties and produce results which are not very reliable even if Blanchard's model is true. Hence, given the present state of the art, it seems to be advisable to look at the Ricardian hypothesis beyond its implications for consumption behavior.

4. Interest Rate Tests

It is therefore interesting to see whether evidence against REQ shows up in time series other than consumption. Interest rates are a suggestive candidate, of course, and quite a bit of empirical work has been devoted to the question whether an increase in debt financing of government absorption actually drives up interest rates. For such tests, see e. g. Plosser (1982), it is typically assumed that the behavior of real interest rates can be described by a vector autoregressive framework in which various exogenous variables exist which may or may not affect real interest rates. It is easy to account for rational expectations and the effects of future developments on current interest rates in such a setup. Since real interest rates are unobservable, nominal interest rates are defined as real rates plus expected inflation, where expected inflation is derived from forecasts of an appropriate autoregressive process. See Evans (1987) for the formal statement of such a model. By substitution, then, a VAR process for nominal interest rates, the inflation rate and exogenous variables is derived and the significance of the exogenous variables can be used to draw inference on the empirical validity of the Ricardian equivalence proposition.

To avoid the curse of dimensionality, I will stick to a rather simple system of four variables: Interest rates, the share of government absorption in GDP (denoted GQ), the share of deficit financing in government expenditures (DEFQ), and the inflation rate of the GDP-deflator (π). As white noise residuals in the VAR can be achieved only in a rather unparsimonious lag-specification, I prefer a single equations approach in which I model both the long-run and the short-run dynamics simulataneously. Denoting the (nominal) long-run interest rate[19] by *ILR*, I estimate the following error-correction equation:

$$\begin{aligned}\Delta ILR_t &= c_1 + c_2 ILR_{t-1} + c_3 GQ_{t-1} + c_4 DEFQ_{t-1} + c_5 \pi_{t-1} \\ &+ c_6 \Delta ILR_{t-1} + c_7 \Delta ILR_{t-4} + c_8 \Delta GQ_{t-1} + c_9 \Delta GQ_{t-4} \\ &+ c_{10} \Delta DEFQ_{t-1} + c_{11} \Delta DEFQ_{t-4} + c_{12} \Delta \pi_{t-1} + c_{13} \Delta \pi_{t-4} \\ &+ c_{14} S1_t + c_{15} S2_t + c_{16} S3_t + c_{17} \varepsilon_{t-1} + c_{18} \varepsilon_{t-4} + \varepsilon_t \end{aligned} \tag{11}$$

Here ε_t is white noise and of course unrelated to the same symbol used in equations (5) to (7). An analogous regression is run with a short-run interest rate *ISR* (three months to maturity). Under REQ one would expect insignificance of the deficit variable in the long run (c_4=0) and in the short run (c_{10}=c_{11}=0). So long as government absorption does not alter the marginal productivity of capital, one would further expect c_3=0 and c_8=c_9=0. Moreover, under long-run monetary neutrality, nominal interest rates and the rate of inflation should cointegrate with cointegrating vector (1, -1)', i. e. we should have c_2=$-c_5$. The results of estimation by OLS and various tests on the significance of the regressors are given in Table 9[20].

[19]This is the so-called "Umlaufrendite".

[20]Again, a standard t-distribution has been used. However, using the critical values of Banerjee, Dolado Mestre (1994) would lead to the same conclusions about significance / insignificance of regressors as does the analysis below.

Table 9

Error Correction Estimates for Interest Rates

Coefficient (Regressor)	Dependent Variable ΔILR		ΔISR	
c_2 (ILR_{t-1})	-0.083 ** (0.026)	-0.028 ** (0.005)	-0.109 ** (0.027)	-0.110 ** (0.019)
c_3 (GQ_{t-1})	0.004 (0.011)	-	-0.023 (0.023)	-
c_4 ($DEFQ_{t-1}$)	0.004 (0.005)	-	0.012 (0.008)	-
c_5 (π_{t-1})	0.088 (0.050)	$-c_2$	0.118 (0.101)	$-c_2$
c_6 (ΔILR_{t-1})	0.126 (0.084)	0.120 * (0.058)	0.300 * (0.116)	0.212 ** (0.061)
c_7 (ΔILR_{t-4})	0.597 ** (0.088)	0.574 ** (0.071)	0.557 ** (0.089)	0.545 ** (0.078)
c_8 (ΔGQ_{t-1})	-0.036 (0.070)	-	-0.286 ** (0.108)	-0.341 ** (0.097)
c_9 (ΔGQ_{t-4})	-0.005 (0.047)	-	-0.003 (0.072)	-
c_{10} ($\Delta DEFQ_{t-1}$)	-0.005 (0.006)	-	0.011 (0.013)	-
c_{11} ($\Delta DEFQ_{t-4}$)	-0.002 (0.005)	-	0.022 (0.013)	-
c_{12} ($\Delta\pi_{t-1}$)	-0.021 (0.023)	-	-0.030 (0.041)	-0.060 (0.016)
c_{13} ($\Delta\pi_{t-4}$)	0.018 (0.028)	-	0.024 (0.040)	-
Q(16)	21.14	24.28	21.84	20.36
$\bar{R}^2$	0.242	0.254	0.433	0.416
Log Likelihood	548.03	547.89	473.42	471.31
*=5% significant, **=1% significant				
Wald-Tests of coefficient restrictions: χ^2-test (P-value)				
H0				
$c_2=-c_5$	0.023 (0.878)		0.012 (0.914)	
$c_3=c_4=0$	0.760 (0.684)		3.336 (0.189)	
$c_3=c_4=c_8=c_9=c_{10}=c_{11}=0$	1.943 (0.925)		17.927 (0.006)	
$c_2=-c_5$ $c_3=c_4=c_8=c_9=c_{10}=c_{11}=0$	2.201 (0.948)		-	
$c_3=c_4=c_{10}=c_{11}=0$	-		6.333 (0.176)	
$c_2=-c_5$ $c_3=c_4=c_{10}=c_{11}=0$	-		8.156 (0.086)	

Looking at the first column of results in Table 9 we see that the only levels variable which is, in fact, significant at conventional significance levels, is the lagged interest rate. The shares of government absorption and the public deficit are clearly insignificant. Also insignificant is the level of inflation rate. Here, however, the associated P-value is 0.08, and the point estimate is quite precisely the negative of the point estimate for the level of the interest rate. I therefore compute a Wald-test of the hypothesis c_2=–c_5, and find that this hypothesis is clearly accepted (see lower panel of Table 9). Similarly, the redundancy of GQ(-1) and DEFQ(-1) is easily confirmed. Thus, for the long-run interest rate there is no long-run evidence against REQ. Rather, it seems that in the long run *ILR* depends solely on the real interest rate.

Turning to the combined short-run and long-run effects, Wald-tests show that GQ and DEFQ do not have any explanatory power for changes in the long run rate (c_3=c_4=c_8=c_9=c_{10}=c_{11}=0). This conclusion is independent of whether or not the restriction c_2=–c_5 is imposed. By succesively deleting insignificant regressors I eventually obtain a specification in which the change in the long run rate depends on no other variable but changes in its own past and the level of the real interest rate (second column of Table 9).

For the short-run rate, results are similar, but not quite alike. In the general regression equation of type (11) the short-run interest rate is the only significant levels variable, cf. column 3. The hypothesis c_3=c_4=0 is accepted and the same is true for the hypothesis c_2=–c_5 despite the fact that the level of the inflation rate is not significant as an isolated regressor. Lagged changes of the levels variables are significant only in the case of the dependent variable's own past and for the lagged share of government absorption (c_8). The sign of this latter coefficient is rather surprising since it implies that more government absorption lowers the short run rate, i. e. there is short run crowding in. Hence, the hypothesis that government absorption and debt do not affect short run interest rates at all is soundly rejected (c_3=c_4=c_8=c_9=c_{10}=c_{11}=0). This, however, does not refute Ricardian equivalence, since the hypothesis that the debt share is irrelevant to short-run interest rates and the share of government absorption does not have long run effects is accepted (c_3=c_4=c_{10}=c_{11}=0) regardless of whether c_2=–c_5 is imposed or not.

Succesively reducing insignificant regressor results in a final specification (column 4) with the additional feature that the lagged change in inflation has a negative impact on short run interest rates. This basically says that in times of rising inflation the nominal interest rate may temporarilly be below its long-run equilibrium level, since e. g. agents have not yet become fully aware of this development. This might be taken as evidence of some type of myopia with respect to monetary policy and its transmission into the real sector of the economy, but it is less clear what this means for Ricardian equivalence. Since imperfect anticipation of inflationary developments might set incentives for the government to inflate a certain percentage of outstanding debt away, such a coefficient could, in principle, be seen as indicative of some non-Ricardian behavior in the short run. However, this sort of evidence is certainly very weak and probably does not constitute a major challenge to Ricardian equivalence: A non-Ricardian effect, if it exists at all, is limited to short-run dynamics; and due to the Lucas-critique, there isn’t, not even in the short-run, a reasonable scope for exploiting this feature in a sizable way.

5. Exchange Rate Tests

One might argue that in an economy with a large open sector such as the German economy, capital mobility will prevent interest rates from responding to increases in the government deficit. Instead, the domestic currency will appreciate and hence the nominal exchange rate will fall. Consequently, one should look for evidence against the Ricardian equivalence proposition on foreign exchange markets rather than in interest rate time series.

Several authors have used this idea to construct econometric tests of REQ with exchange rate data, see e. g. Evans (1986) and Feldstein (1986). In this paper, I follow Beck's (1994) synthesis of earlier work. Beck essentially derives two alternative equations which might be useful in describing the change in spot exchange rates. Suppressing constants, dummies and error terms for notational convenience, the two equations postulate

$$\Delta e_t = a_1\left(\mu_t - \mu_t^e\right) + a_2\left(\pi_t - \pi_t^e\right) + a_3\left(d_t - d_t^e\right) + a_4\left(g_t - g_t^e\right) \tag{12}$$

or

$$\Delta e_t = b_1\left(\mu_{t+1}^e - \mu_t^e\right) + b_2\left(\pi_{t+1}^e - \pi_t^e\right) + b_3\left(d_{t+1}^e - d_t^e\right) + b_4\left(g_{t+1}^e - g_t^e\right), \tag{13}$$

where Δe_t is the change in the expected rate of return of a foreign investor, μ_t is the rate of growth of the money supply, d_t is the share of deficit financing in GDP and g_t is the share of net government absorption in GDP. A superscript e denotes the expectation of the variable formed one period earlier, i. e. μ_t^e is *t-1*'s expectation of μ_t. To make my description of e_t more precise, I have constructed e_t from

$$e_t = ISR_t + \ln s_t - \ln f_t$$

where s_t is the DM/$ spot rate and f_t is the corresponding forward rate. I have used the three-months interest rate since a period in this application is a quarter. For the same reason, I use a narrow money definition (M1) for computing μ_t.

In order to obtain time series for the expected variables, I estimate a vector error correction model (VECM) for the four variables μ_t, π_t, d_t, and g_t and compute the one-step-ahead forecasts. I find evidence for one cointegrating vector yielding the stationary combination g_t-0.96 d_t+3.15 μ_t-3.37π_t, which looks quite reasonable since it suggests that the pairs g_t and d_t as well as μ_t and π_t have equal coefficients in absolute value, but differ in sign[21]. Estimating the VECM with four lags in differences yields white noise residuals and from these I obtain the one-step ahead forecasts that replace the expected variables in equations (12) and (13).

Under the null of REQ the coefficients a_3 and b_3 should be zero, while under the alternative of conventional open macroeconomics they should be positive. Neither hypothesis makes clear predictions about the signs of the coefficients a_4 and b_4, though. In estimating (12) and (13) it is clear that there is a simultaneity problem, as the regressors are likely to be correlated with the error term. It is, however, very difficult to find suitable instruments (I have been unable to find any that are reasonably correlated with the regressors they are intended to replace). As bad instruments may make things actually worse than they are under OLS (cf. Nelson and

[21] I have not explored this issue any further, since it is not central to the following analysis. The coefficients of the cointegrating vector are highly significant by conventional levels of significance.

Startz (1990)), I have estimated (12) and (13) by OLS despite the simultaneity problem. Another justification for that is given below.

Table 10

Exchange Rate Tests

a_1 $(\mu_t\text{-}\mu_t^e)$	-0.089 (0.053)	b_1 $(\mu_{t+1}{}^e\text{-}\mu_t^e)$	-0.090 (0.046)
a_2 $(\pi_t\text{-}\pi_t^e)$	0.127 (0.135)	b_2 $(\pi_{t+1}{}^e\text{-}\pi_t^e)$	-0.103 ** (0.040)
a_3 $(d_t\text{-}d_t^e)$	0.081 (0.065)	b_3 $(d_{t+1}{}^e\text{-}d_t^e)$	0.042 (0.041)
a_4 $(g_t\text{-}g_t^e)$	0.148 (0.272)	b_4 $(g_{t+1}{}^e\text{-}g_t^e)$	-0.210 (0.187)
$\bar{R}^2$	0.100	$\bar{R}^2$	0.148
*=5% significant, **=1% significant			

From Table 10 we see that the deficit variable is not significant in either specification. Hence it seems difficult to argue that the failure of interest rate based tests for REQ to provide evidence against the null is due to open economy considerations: A significant impact of deficits on exchange rates is not conceivable. The only concern that remains is thus the fact that the use of OLS in these regression might lead to misguided inference. But one can easily check how well-founded this concern is. By estimating a VECM for the variables e_t, μ_t, π_t, d_t, and $g_t$we get estimates of the covariance matrix of contemporaneous innovations. According to these estimates none of the covariances is greater than 0.25 in absolute value, most of them, are below 0.1. One might hence conjecture, that the simultaneity bias is not substantial and the results given in Table 10 are by and large reliable.

6. Conclusions

In this paper, I have worked through the most prominent tests of Ricardian equivalence and used them with German data. Interestingly, there is hardly any evidence against REQ. Given this result, it is quite surprising that the theoretical model which states REQ in the presence of a fully rational, optimizing economic agent is strongly rejected by the data. Hence, it seems that future research should address the issue of modifying the representative consumer model to allow for alternative utility functions, non-separabilities, habit persistence or the like to find a specification, where rational optimizing behavior on the part of the agent implies Euler-equations that actually fit the data. Once this is done, we might be able to perform powerful tests of REQ using Blanchard's (1985) model of finite horizons. If, however, one were unable to find a useful model for the aggregate consumer's behavior, REQ would constitute an important empirical puzzle, as it could well be traced in the data but would lack a sound theoretical foundation.

Data Appendix

All data are quarterly data for West Germany (1960.1-1994.4) and not seasonally adjusted. (After reunification, a separate system of national accounts has been maintained for West Germany until 1994.4). All flow data are taken from the Deutsches Institut für Wirtschaftsforschung (DIW), Berlin. (This data set also contains the data on interest rates, exchange rates and terms of trade). Stock data for monetary aggregates were derived by using annual values on monetary wealth from the Deutsche Bundesbank and interpolating quarterly observations by using the corresponding flows from the DIW's system of national accounts. Since the Bundesbanks stock data for West Germany do not extend beyond 1992, quarterly values for the last two years have been derived by simply accumulating the DIW's West German flow data.

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Abstract

The approximate validity of the Ricardian equivalence proposition remains hotly disputed despite a large body of empirical work devoted to this issue. Contradictory conclusions arise at least partially from the fact that a large variety of very different tests has been applied to very different data sets. Moreover, instationarity properties of the data are not properly taken into account in some of the older literature, so that inference may be misguided. This paper collects the most important tests of Ricardian equivalence and applies them to a single data set of quarterly West German macro data. It turns out that there is hardly any evidence against the Ricardian proposition. Paradoxically, the theoretical model that implies Ricardian equivalence, is strongly rejected by the data.

Zusammenfassung

Trotz einer großen Anzahl empirischer Arbeiten ist es seit längerem heftig umstritten, ob die Ricardianische Äquivalenzhypothese als approximativ zutreffend angesehen werden kann. Widersprüchliche Ergebnisse derartiger Studien sind zumindest partiell darauf zurückzuführen, daß unterschiedliche Tests mit unterschiedlichen Datensätzen durchgeführt wurden. Auch wurden die Instationaritätseigenschaften der Daten in einem Teil der älteren Literatur nicht ausreichend gewürdigt, so daß deren Schlußfolgerungen u. U. fehlerbehaftet sind. In diesem Aufsatz werden die wichtigsten Tests der Ricardianischen Äquivalenzhypothese auf einen einheitlichen Satz westdeutscher Makrodaten angewandt. Erstaunlicherweise findet sich fast keine empirische Evidenz gegen die Äquivalenzhypothese, das theoretische Modell jedoch, das ihr zugrundeliegt, wird von den Daten klar verworfen.