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Re-examining the Decline in the US Saving Rate: The Impact of Mortgage Equity Withdrawal*

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Abstract

In this paper we examine the role of mortgage equity withdrawal in explaining the decline of the US saving rate, since when house prices rise and mortgage rates are low, homeowners have an incentive to withdraw housing equity and this may affect the saving rate. We estimate a Vector Error Correction (VEC) model including the saving rate, asset prices, equity withdrawal and interest rates and find that indeed mortgage equity withdrawal is a key determinant of the observed saving pattern.

Keywords: Saving rate, Mortgage equity withdrawal, Asset prices, Mortgage rates, Vector Error Correction, Impulse response analysis
JEL Classification: C32, E21, O51

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1 Introduction

The US saving rate has been declining sharply since the 1990s. The personal saving rate dropped from an average of 8.6 percent in the period 1980-1990 to an average of 5.5 in the period 1990-2000. The average rate has fallen to 3.5 percent over the period 2000-2011. This decline is now considered a stylized fact and has attracted a lot of attention from academics and policy-makers. Greenwood and Jovanovic (1999) put forward the idea that recent advances in technology and in labour productivity have led US households to revise upwards their permanent income estimates. Lusardi et al. (2001) take the view that the appreciation of assets and the increase in medical care expenditure are the causes of the drop in the personal saving rate. Muellbauer (2008) argues that significant improvements in credit access have increased the ability of households to extract or borrow against their home equity changing the saving behaviour in many countries.

In this paper we focus on mortgage equity withdrawal (MEW - also known as home equity extraction) as a possibly important cause of the decline in the US saving rate. This is defined as the amount of equity that is extracted from an asset when it appreciates. In general, when house prices are rising and mortgage rates are low, homeowners have an incentive to with-draw housing equity (see e.g. Duca and Kumar, 2011; Paradiso et al., 2012; Paradiso, 2012) and this may increase consumption expenditure. Smith and Searle (2008) argue that in the last few years MEW has been the main trans-

mission mechanism for housing wealth effects onto the aggregate economy; Greenspan and Kennedy (2005) and Hatzius (2006) also take the view that MEW has played a crucial role in determining private consumption expenditure. Empirical studies for the US show that regressions of consumption on mortgage equity withdrawal yield coefficients ranging from zero to as high as 0.62 for the long-run propensity to consume (Catte et al., 2004; Hatzius, 2006; Klyuev and Mills, 2007; Girouard, 2010). Catte et al. (2004) find that MEW drives consumption with a marginal propensity to consume equal to 0.2 for the US when an error correction model including consumption, disposable income, net financial wealth, net housing wealth and MEW variables is estimated. Using a single equation error correction model, Hatzius (2006) finds that each dollar of MEW generates 62 cents of extra consumer spending when the consumption ratio, net wealth, interest rate and MEW are taken into account. Klyuev and Mills (2007) study the role of MEW in explaining the decline in the saving rate for different countries. Their empirical results for the US indicate that MEW is not statistically significant in a single equation error correction model with the saving rate, net wealth, interest rates and inflation. Girouard (2010) investigates the effects of housing wealth on the marginal propensity to consume in the US and other OECD countries and shows that they are stronger where mortgage markets are "most complete", in particular where they provide opportunities for MEW.

This paper aims to contribute to the current literature on the decline of the US saving rate over the period 1990-2011 by focusing on the role

of MEW in a multivariate time series framework. Specifically, the analysis improves on the earlier studies discussed above in two respects. First, a VEC model is estimated instead of a single equation error correction model. This is important since the assumption of exogeneity implicitly made in a single equation model for the right-hand side variables (see Urbain, 1992; Ericsson and MacKinnon, 2002) may not be a valid one for MEW and house prices (see Mishkin, 2007; André et al., 2011, among others). By contrast, in the Johansen (1988) approach used here all variables are jointly modelled in a complete closed form model, full information analysis can be carried out and the number of co-integrating vectors can be determined performing appropriate co-integration tests. Second, the estimation of a multivariate model instead of a single equation one allows to investigate the dynamic linkages between the variables using impulse response analysis, a valuable tool in cointegrated systems (see Lütkepohl and Reimers, 1992). Since in such a system the deviations from equilibrium are stationary, they will eventually revert to equilibrium, and their time paths provide useful insights into the short-run and long-run relationships between the variables of the system. Therefore, the chosen framework enables us to investigate both the short- and the long-run impact of mortgage equity withdrawal on the saving rate.

In the empirical analysis, we first estimate a VEC model with five variables typically considered in the literature (see i.e. Hatzius, 2006; Klyuev and Mills, 2007), namely the saving rate, the stock price index, house prices, mortgage interest rates and MEW. Because the signs of the estimated long-

4

run coefficients on house prices and mortgage interest rates are not consistent with economic theory we test the restrictions that both these cointegration coefficients and the corresponding factor loadings are zero. Since these restrictions are found to hold, we then proceed to estimate a three-variate VEC model without house prices and interest rates. A significant long-run relationship is found between the remaining variables, and the impulse-response analysis shows that mortgage equity withdrawal indeed drives the saving rate.[1]

The paper is organised as follows. Section 2 describes the data. Section 3 presents the empirical results. Section 4 offers some concluding remarks.

2 Data description

The data used for the empirical analysis cover the period 1993:Q1-20011:Q1. The series are: the saving rate, the stock market index, house prices, the nominal mortgage rate and mortgage equity withdrawal. The saving rate is the personal saving rate and the data have been obtained from the Bureau of Economic Analysis (BEA). The stock market index and housing prices are the Standard and Poor's 500 index in logarithms and the year-on-year growth rate of the Standard and Poor's/Case-Shiller home price index, respectively. Both have been deflated using the US consumer price index (CPI). The CPI series is from the Federal Reserve Economic Data (FRED) database

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[1]For another study on the real economy, house prices and mortgage rates see Rubio, 2011.

maintained at the Federal Reserve Bank of St. Louis.

Although other studies (see, e.g., Klyuev and Mills, 2007) use net wealth variables, we choose instead stock and house prices as asset prices. The reason is that stock values may affect spending either through wealth effects or through their role as leading indicators of income and job growth (see Poterba and Samwick, 1995). In addition, stock price fluctuations may influence consumption by affecting consumer confidence. A similar reasoning applies also to house price changes. Therefore, we focus on asset prices as variables containing relevant information for explaining the decline in the US saving rate.[2] In addition, we include the nominal mortgage interest rate. Several other studies also consider the nominal interest rate as an additional variable in the US consumption and saving functions (Mishkin, 1976; Gylfason, 1981; Wilcox, 1990; Klyuev and Mills, 2007, among others), since low interest rates are thought to have led to higher personal borrowing and to have fuelled the consumer boom over the last 20 years (Chen and Winter, 2011). The nominal mortgage interest rate is used here for two reasons. First, the increase in household debt in recent years can mostly be attributed to the huge increase in home-related mortgage debt and, to a lesser extent, to pure consumer credit.[3] Second, the recent innovations in the mortgage mar-

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[2]In a recent paper, Chauvin et al. (2011) also emphasise the role of asset prices in explaining consumption.

[3]Mortgage debt increased from about 60% of disposable income in the 1990s to about 83% in the early years of this century, whilst consumer debt rose from about 17% of disposable income in 1960 to only 25%. Home mortgages and consumer debt represented 74% and 22% respectively of the nearly 6 trillion dollar increase in household debt between 1990 and the early 2000s (for further details, see the Federal Reserve Board (FRB) and

6

ket have reduced transactions costs and increased cash-out refinancing (see Cynamon and Fazzari, 2008).

MEW is the equity extracted from the existing homes via cash-out refinancing, home equity borrowing and housing turn-over (see Greenspan and Kennedy, 2008). Specifically, "active" MEW consists of the cash-out refinancing and home equity borrowing that are discretionary actions to extract home equity while "passive" MEW is the equity released during housing turn-over. In our analysis we consider "active" MEW, expressed as a ratio to disposable income, since the literature on the saving-consumption ratio has shown that active MEW has strongly affected consumption. In particular, a survey of the Federal Reserve conducted during the years 2001-2002 shows that consumers used 16% of the equity extracted through cash-out refinancing for consumer expenditure and 35% for home improvements, while they used the remainder to repay other debts, to make other investments or to pay taxes (see Canner et al., 2002). The data are taken from the Greenspan and Kennedy's (2008) data set.[4]

3 Empirical results

In this section, we present the empirical analysis based on a VEC model. As preliminary step, we investigate the unit root properties of the variables using

Bureau of Economic analysis (BEA), and Kim, 2011).

[4]We are grateful to Greenspan and Kennedy who provided an updated series of active MEW (1993:Q1-2011:Q1). The series is not seasonally adjusted. We have carried out the seasonal adjustment with X-12 ARIMA using the Demetra package.

the ADF and DF-GLS tests. The results are reported in Table 1. The null hypothesis of a unit root cannot be rejected for the levels of all five variables. We also test for the null of a unit root in the first differences, which can be rejected at the 1% significance level, with the exception of ADF test for *mew* (5%). Overall, the evidence from the ADF and DF-GLS tests clearly indicates that all variables can be characterised as a unit root process. Since all series are I(1), it is legitimate to test for cointegration. Therefore we estimate an unrestricted VAR that forms the basis for system cointegration tests (see Lütkepohl, 2004).

The VAR model includes the saving rate (sr), the stock price index ($sp500$), the house price index (hp), the mortgage interest rate ($imor$) and mortgage equity withdrawal (mew). In order to select the lag length of the VAR several information criteria are considered. The FP, SIC and HQ criteria suggest a VAR model with two lags. A series of diagnostic tests for the VAR specification with the chosen number of lags are reported in Table 2. In particular, we test for autocorrelation and non-normality in the VAR(2) residuals. The results are satisfactory with the exception of a suggestion of non-normality (see Table 3). An absolute value of unity or less for skewness is acceptable according to Juselius (2006). Furthermore, since Johansen's (1988) multivariate approach appears to be robust to excess kurtosis, non-normality does not seem to be a serious problem (see Juselius, 2001).

After checking for the adequacy of a VAR(2) specification, we proceed to test for cointegration using the trace test proposed by Johansen (1988).

Table 1: Unit root test results.

Variable	ADF	DF-GLS
sr	-2.614^{*}	-1.263
Δsr	-12.049^{***}	-10.093^{***}
$sp500$	-2.206	-0.912
$\Delta sp500$	-5.067^{***}	-5.100^{***}
hp	-1.281	-1.269
Δhp	-5.048^{***}	-4.015^{***}
$imor$	-1.199	-0.402
$\Delta imor$	-6.882^{***}	-4.435^{***}
mew	-0.767	-0.821
Δmew	-3.086^{**}	-3.062^{***}

Notes: ***, ** and * denote significance at the 1%, 5% and 10% level respectively. A model with a constant is considered. The number of lags for the ADF and DF-GLS tests is selected according to the Schwert (1989) information criterion. The critical values for the ADF and the DF-GLS unit root tests are tabulated in MacKinnon (1996) and Elliot et al. (1996) respectively.

The results are reported in Table 4. They show that the null of rank $r = 1$ cointegrating vectors cannot be rejected at the conventional significance levels.

Table 2: Diagnostic tests for the VAR(p) specification. *sr*, *sp500*, *hp*, *imor* and *mew* variables.

p	Q_{16}	FLM_5	LJB_5^L	$MARCH_{LM}(4)$
2	376.553 [0.16]	1.311 [0.06]	18.674 [0.00]	893.327 [0.56]

Notes: p-values are in parenthesis. Q_h indicates the multivariate Ljiung-Box Portmentau test. FLM_h is a variant of the Breusch-Godfrey LM test for autocorrelation up to order h. LJB_K^L is the multivariate Lomnicki-Jarque-Bera test for non-normality; $MARCH_{LM}(q)$ is the multivariate LM test for ARCH.

Table 3: Univariate tests for normality. *sr*, *sp500*, *hp*, *imor* and *mew* variables.

tests	$sp500$	hp	$imor$	mew	sr
Norm(2)	0.651 [0.72]	45.831 [0.00]	11.222 [0.00]	7.780 [0.02]	1.422 [0.49]
Skewness	-0.018	-0.90	0.58	-0.47	0.29
Excess Kurtosis	2.70	6.50	4.56	4.33	3.39

Notes: p-values in brackets.

Therefore, for the VECM analysis we assume a single cointegrating vector. The VECM estimation results are reported in Table 5. They show that the estimated coefficients for both *hp* and *imor* have a sign contradicting theory. As pointed out by Duca and Kumar (2011), mortgage equity withdrawal should be encouraged by a decrease in long-term interest rates and an increase in house prices. This suggests that *mew* may capture part of the information embodied in *hp* and *imor*, and this may explain the wrong signs of these

Table 4: Cointegration results. *sr*, *sp500*, *hp*, *imor* and *mew* variables

$H_0 : r =$	Trace Statistics	CV10%	CV5%	CV1%
r = 0	89.00	72.74	76.81	84.84
r = 1	48.92	50.50	53.94	60.81
r = 2	21.91	32.25	35.07	40.78
r = 3	9.44	17.98	20.16	24.69
r = 4	3.73	7.60	9.14	12.53

Notes: Sample 1993:Q1-20011:Q1. *r* indicates the number of cointegrating vectors. Deterministic terms in the model: constants and three spike dummies (2002:Q3, 2004:Q2, 2008Q4). The first dummy is included for the stock market crash, the second one for the sharp rise in the house prices and the third one for economic recession (GDP fell by 7%). The dummies are not restricted to the long-run. The critical values of Johansen's trace tests are obtained by computing the relevant response surface as in Doornik (1998).

variables in the cointegrating vector. When house prices increase or interest rates decrease, consumers are willing to withdraw cash from housing equity by increasing their mortgage loans or re-negotiating their mortgage contracts at better conditions. These additional funds can be used for a variety of purposes such as consumption. As regards the other two variables, *mew* and *sp500*, the results indicate a negative effect on the saving rate. A rise in *sp500* increases the value of financial assets owned by households, i.e. their financial wealth, and therefore consumption through wealth effects. The coefficient on *mew* has a negative sign as expected. This may be due to the fact that in the last twenty years MEW has become an important resource for consumer spending, being favoured by financial innovations (which lowered the costs of withdrawing housing equity via cash-out mortgage refinancing), an increase in house prices and lower interest rates (see Greenspan and Kennedy, 2005;

Hatzius, 2006; Paradiso et al., 2012). The sign of the speed of adjustment coefficient on *sr* is negative, but the coefficients on *imor* and *hp* have the wrong signs, suggesting that they contain only redundant information already embodied in *mew*.

Table 5: Cointegration vector and loading parameter for VECM and cointegrating rank $r = 1$. *sr*, *sp500*, *hp*, *imor* and *mew* variables.

	sp500	mew	hp	imor	sr	cons
$\hat{\beta}'$	1.858 (4.48)	0.683 (6.68)	−0.107 (−3.91)	0.101 (0.86)	1	−8.917 (−7.34)
$\hat{\alpha}'$	0.025 (3.26)	−0.201 (−3.58)	0.596 (2.82)	−0.057 (−1.14)	−0.365 (−3.25)	

Notes: Sample 1993:Q1-20011:Q1. t-statistics in parentheses.

Since the estimated coefficients on *hp* and *imor* are inconsistent with theory, we proceed to test whether the model can be reduced, that is if some valid restrictions can be imposed. In particular, theory suggests that the *mew*, *imor* and *hp* variables should convey the same information because house equity is extracted when house prices and interest rates are low. This can be tested by imposing restrictions on both the cointegrating vector and the factor loading coefficients. Specifically, we test whether the coefficients of *hp* and *imor* in the cointegrating vector and their factor loadings are all zero. The results show that the restrictions hold (the p-value is equal to 0.17). For this reason, we proceed to estimate a VAR model with only three variables, *sr*, *mew* and *sp500*. The information criteria suggest different lag orders (specifically AIC, HQ and FPE suggest three lags, whilst SIC only one). We opt for three lags on the basis of the diagnostic tests (see Table 6).

12

Table 6: Diagnostic tests for VAR(p) specification. *sr, mew* and *sp500* variables.

p	Q_{16}	FLM_5	LJB_3^L	$MARCH_{LM}(5)$
3	110.96 [0.64]	1.371 [0.09]	28.163 [0.00]	195.67 [0.20]

Notes: p-values are in parenthesis. Q_h indicates the multivariate Ljiung-Box Portmentau test. FLM_h is a variant of the Breusch-Godfrey LM test for autocorrelation up to order h. LJB_K^L is the multivariate Lomnicki-Jarque-Bera test for non-normality; $MARCH_{LM}(q)$ is the multivariate LM test for ARCH.

As for the previous VAR specification, we check for non-normality in the residuals (see Table 7). Again, non-normality is a not a problem since it appears to be caused by excess kurtosis (see Juselius, 2001, 2006).

Table 7: Univariate tests for normality. *sr, sp500* and *mew* variables.

tests	$sp500$	mew	sr
Norm(2)	15.386 [0.00]	4.054 [0.14]	0.601 [0.74]
Skewness	-0.607	-0.084	0.223
Excess Kurtosis	4.942	4.167	3.019

Notes: p-values in brackets.

Therefore we test for cointegration. The results are reported in Table 8. They imply that the null of one cointegrating vector cannot be rejected at conventional significance levels.

We then estimate the cointegrating vector. The results are reported in Table 9. All the variables have the expected signs and the factor loading for *sr* is negative and statistically significant, which implies that the long-run relationship for *sr* holds; this supports the interpretation of the cointegrating

13

Table 8: Cointegration results. *sr*, *mew* and *sp500* variables.

$H_0 : r =$	Trace Statistics	CV10%	CV5%	CV1%
r = 0	38.86	32.25	35.07	40.78
r = 1	10.40	17.98	20.16	24.69
r = 2	3.80	7.60	9.14	12.53

Notes: Sample 1993:Q1-2011:Q1. r indicates the number of cointegrating vectors. Deterministic terms in the model: constants and three spike dummies (2001:Q2, 2003:Q3 and 2004:Q2). The first dummy variable is included for the boom in the house prices (see Dreger and Kholodilin, 2011), and the second for the boom in the mortgage refinancing. For the third dummy variable, see the notes to the Table 4. The dummies are not restricted to the long-run. The critical values of the Johansen's trace tests are obtained by computing the relevant response surface according to Doornik (1998).

vector as a stationary saving function with the saving rate being related to *mew* and the stock market index *sp500*. Deviations from the long-run equilibrium are absorbed in less than two quarters (the coefficient is equal to 0.7).

Table 9: Cointegration vector and loading parameter for VECM and cointegrating rank $r = 1$. *sr*, *mew* and *sp500* variables.

	sp500	mew	sr	cons
$\hat{\beta}'$	1.459 $_{(4.09)}$	0.481 $_{(7.54)}$	1	-7.325 $_{(-12.99)}$
$\hat{\alpha}'$	-0.041 $_{(-0.49)}$	0.031 $_{(2.19)}$	-0.713 $_{(-4.82)}$	

Notes: sample 1993:Q1-2011:Q1. t-statistics in parentheses.

Next, we carry out the impulse response function analysis. Within the VEC framework, we use the Cholesky identification strategy and assume the following order of the variables: *sp500*, *mew* and *sr*.[5]

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[5]The Cholesky decomposition is widely used in the empirical literature to identify

A positive shock to *sp500* decreases *sr* in line with the predictions of consumption theory. However, the estimates indicate that *sr* does not react significantly to a *sp500* shock. A positive shock to *mew* decreases *sr*, although the response is statistically significant only after 2 quarters. The response of *sp500* to a *sr* shock is relatively smooth. This is in line with the evidence on the recent stock market bubble with equity prices not following fundamentals. As expected, a rise in *sr* leads to a reduction in mortgage equity withdrawal because households are less willing to extract cash from housing equity, but this effect is not statistically significant.

On the whole, the impulse response analysis suggests that *mew* is the main driving force of the saving rate. A sharp housing appreciation over the last two decades has turned housing into a major store of wealth (see Smith, 2006) and this housing wealth effect has been stronger than other financial wealth effects (Benjamin et al., 2004; Leonard, 2010). Furthermore, as a result of international deregulation, homeowners have renegotiated their mortgage loan contracts. These developments, together with increased borrowing resulting from low interest rates, have decreased the saving rate.[6]

Finally, we test for the stability of the estimated system. Hansen and Johansen (1999) have proposed recursive statistics for stability analysis in

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structural shocks (see Rossi and Zubairy, 2011; Davis and Zhu, 2011; Gibson et al., 2012). Because the results can be sensitive to the order of the variables, Sims (1981) recommends checking their robustness to a different order - this is found to hold in our case.

[6]McCarthy and Peach (2004) note that mortgage interest rates have responded quickly to the monetary policy change after the 1990s. For a study on monetary policy over a longer horizon see Castelnuovo (2012).

the context of a VEC model with cointegrated variables. Because the cointegrating rank is r=1, there is one non-zero eigenvalue. Its confidence intervals and the tau statistics $\tau_T^{(t)}(\xi_1)$ are plotted in Figure 2 together with the critical values for a 5% level test. The recursive eigenvalue appears to be fairly stable, and the values of $\tau_T^{(t)}(\xi_1)$ are considerably smaller than the critical values. Thus, stability of the system appears to be confirmed.

4   Conclusions

This paper contributes to the current literature on the behaviour of the US saving rate by focusing on the role of mortgage equity withdrawal. Whilst previous studies have analysed the relationship between the saving rate, mortgage equity withdrawal, asset prices and interest rates in a single equation error correction model, the present one estimates a Vector Error Correction model since the assumption of exogeneity implicitly made for the right-hand side variables of a single equation model may not be a valid one for mortgage equity withdrawal and house prices. Having estimated a cointegrated system as in Johansen (1988), we examine both the long-run equilibrium by performing cointegration tests and the short-run dynamic linkages by means of impulse response analysis. The initial VEC specification includes the saving rate, the stock price index, house prices, mortgage interest rates and MEW. However, because the signs of the estimated long-run coefficients on house prices and mortgage interest rates are inconsistent with theory, we test for the relevant zero restrictions on both cointegrating coefficients and factor loadings. Since these are found to hold, we then proceed to estimate a three-variate VEC model dropping house prices and interest rates.

We find a significant long-run relationship between the saving rate, stock prices and mortgage equity withdrawal, and all the variables have the expected signs. These results are in contrast to those obtained in previous studies where mortgage equity withdrawal is not significant (see Belsky and

Prakken, 2004; Klyuev and Mills, 2007). In addition, the impulse response analysis shows tha this variable is the main determinant of the dynamics of the saving rate. This has important policy implications for monetary policy, namely it suggests that monetary authorities can influence consumption through this channel in addition to the traditional interest rate and asset prices/wealth effect ones. Changes in interest rates affect consumption not only through the after-mortgage-payments household disposable income (the interest rate channel) and the asset price (the asset prices/wealth effect), but also through housing equity extraction (Paradiso, 2012). This implies that the effect of monetary policy on private spending is amplified through its impact on the equity extraction mechanism. Therefore, monetary policy appears to have an additional channel for controlling demand and output in the economy.

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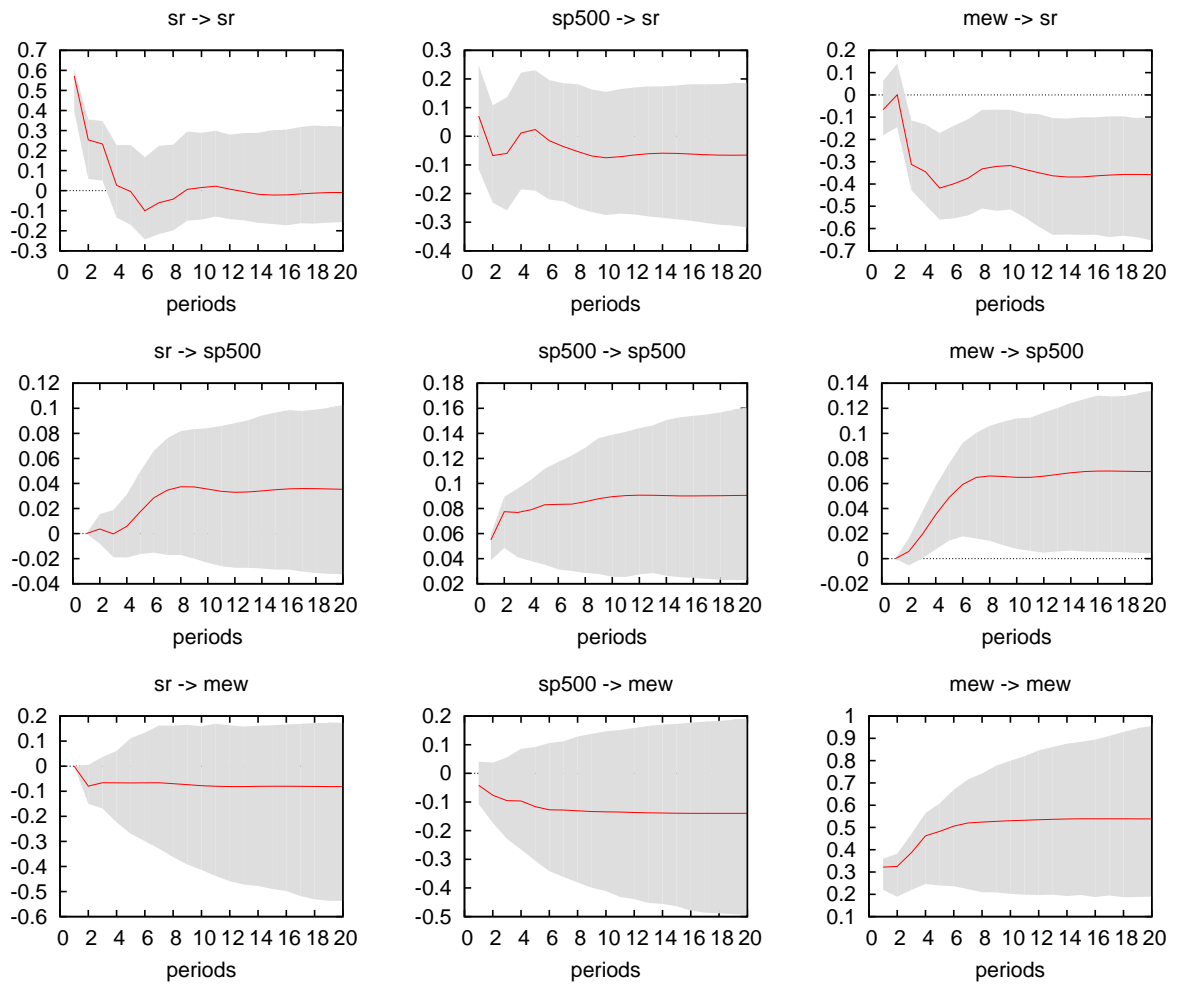Wilcox, J. A., 1990. Nominal interest rate effects on real consumption expenditure. Business Economics, 25, 31-38.

Figure 1: Impulse response analysis for a VEC model with *sr*, *mew* and *sp500* variables with 95% Hall bootstrap confidence intervals based on 2,000 bootstrap replications.
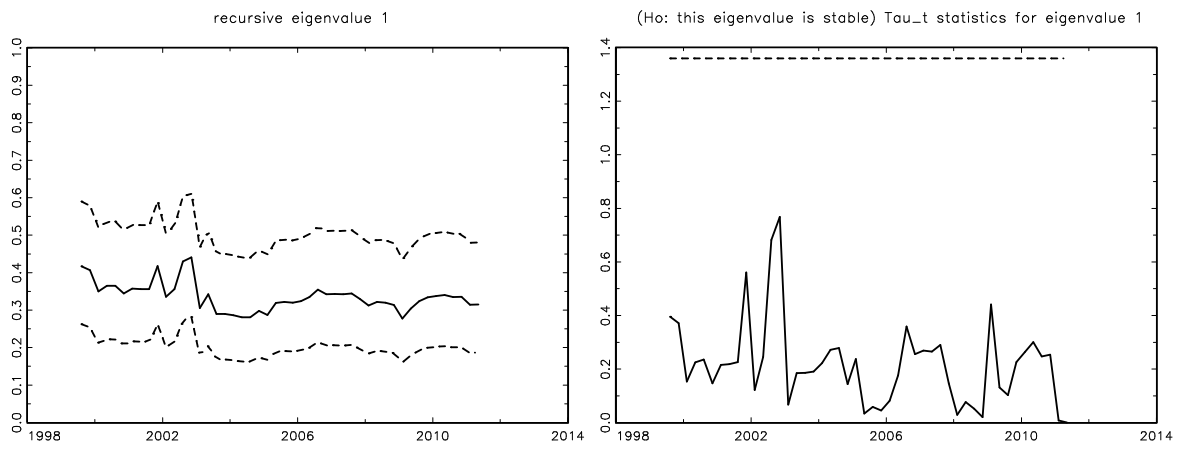
Figure 2: Recursive eigenvalue analysis of VEC model with *sr, mew* and *sp500*. Critical values for a 5% test level.