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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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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. Speci¯cally, 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; Andr¶ 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 to38560(eilibriut11,)-928(and)8561(ti(er)8478(time)856pathsme)856pmprosideu6(fme)8560(i(rig run coe±cients on house prices and mortgage interest rates are not consistent with economic theory we test the restrictions that both these cointegration coe±cients 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 signi¯cant 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 o®ers 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 de°ated using the US consumer price index (CPI). The CPI series is from the Federal Reserve Economic Data (FRED) database

[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 a®ect spending either through wealth e®ects or through their role as leading indicators of income and job growth (see Poterba and Samwick, 1995). In addition, stock price °uctuations may in-°uence consumption by a®ecting consumer con¯dence. 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-

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

[3] Mortgage debt increased from abouoB2MishkinteressD[(Ireased)430(expgd)430(]TJ.423(pap)--391(b)set)3(p

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

MEW is the equity extracted from the existing homes via cash-out re¯-nancing, home equity borrowing and housing turn-over (see Greenspan and Kennedy, 2008). Speci¯cally, \active" MEW consists of the cash-out re¯-nancing 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 a®ected 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 re¯nanc-ing 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

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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 signi¯cance levels.

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

p	Q_{16}	FLM_5	LJB_5^L	$MARCH_{LM}(4)$
2	376:553	1:311	18:674	893:327
	[0.16]	[0.06]	[0.00]	[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.

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311 [0:06]

tests	*sp*500	*hp*	*imor*	*mew*	*sr*

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

$H_0 : r$

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*
β'	1.858 (4.48)	0.683 (6.68)	-0.107 (-3.91)	0.101 (0.86)	1	-8.917 (-7.34)
α'	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 becausTJ/F11-l493111-l49

Table 6: Diagnostic tests for VAR(p) speci¯cation. *sr*, *mew* and *sp500*

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*
β'	1.459 (4.09)	0.481 (7.54)	1	-7.325 ($-$12.99)
α'	-0.041 ($-$0.49)	0.031 (2.19)	-0.713 ($-$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]

[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 signi¯cantly to a *sp500* shock. A positive shock to *mew* decreases *sr*, although the response is statistically signi¯cant 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 e®ect is not statistically signi¯cant.

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 e®ect has been stronger than other ¯nancial wealth e®ects (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 the context of a VEC model with cointegrated variables. Because the cointegrating rank is r=1, there is one non-zero eigenvalue. Its con¯dence intervals and the tau statistics $¿_T^{(t)}(»_1)$ 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 $¿_T^{(t)}(»_1)$ are considerably smaller than the critical values. Thus, stability of the system appears to be con¯rmed.

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 speci¯cation includes the saving rate, the stock price index, house prices, mortgage interest rates and MEW. However, because the signs of the estimated long-run coe±cients on house prices and mortgage interest rates are inconsistent with theory, we test for the relevant zero restrictions on both cointegrating coe±cients and factor loadings. Since these are found to hold, we then proceed to estimate a three-variate VEC model dropping b dropqua-D7(cr)-346(v6.1.9n)-.7e

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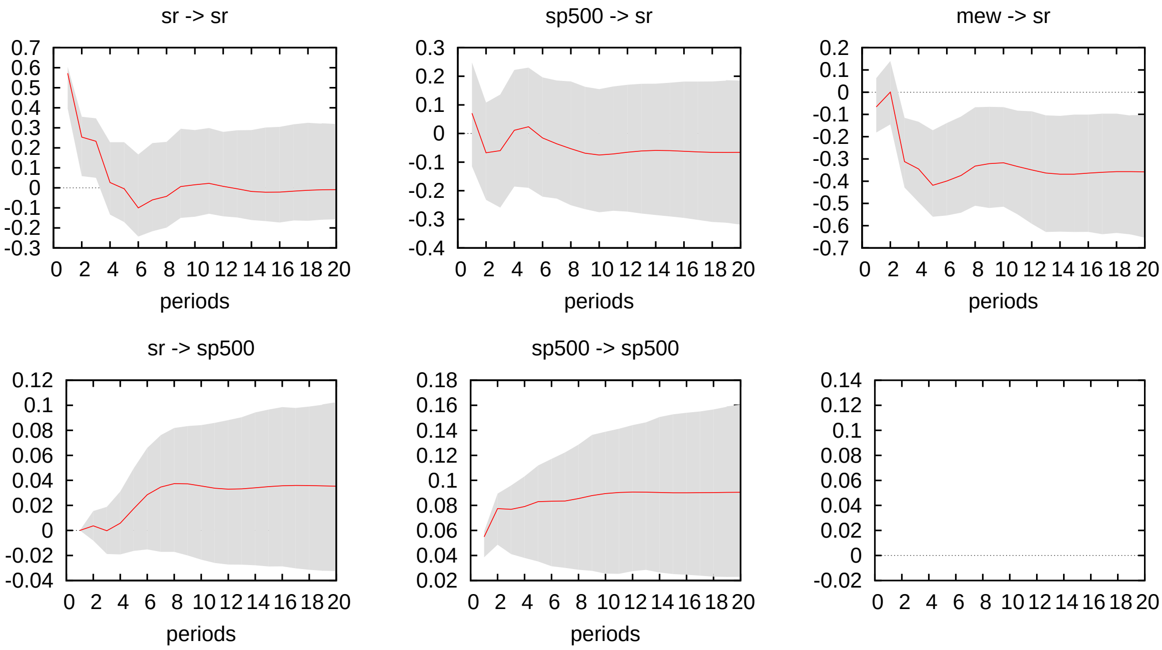

Figure 1: Impulse response analysis for a VEC model with *sr*, *mew* and *sp500* variables with 95% Hall bootstrap con¯dence intervals based on 2,000 bootstrap replications.

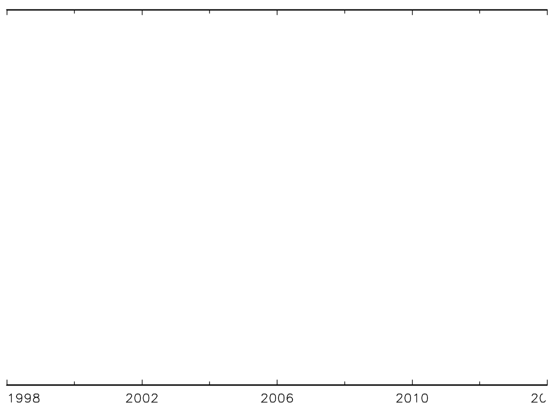

Figure 2: Recursive eigenvalue analysis of VEC model with *sr*, *mew* and *sp500*. Critical values for a 5% test level.