The time-series linkages between US fiscal policy and asset prices*

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Abstract

This article studies the interplay of fiscal policy and asset price returns of the United States in a time-varying-parameter vector autoregressive model. Using annual data from 1890 to 2013, we study the effects of dynamic shocks to both fiscal policy and asset returns on asset returns and fiscal policy. Distinguishing between low volatility (bull market) and high volatility (bear market) regimes together with a time-varying-parameter vector autoregressive model enables us to isolate the different size and sign of responses to shocks during different time periods. The results indicate that increases in the primary surplus to GDP ratio decrease house returns over the entire sample and at each impulse horizon. Unlike the house return response, stock returns only decrease in the first year after the fiscal shock, but then increase for the following eight years. Furthermore, the findings show that asset return movements affect fiscal policy, whereby fiscal policy responds more to equity returns than to house returns. The response of fiscal policy to asset returns proves relatively stable and constant over time while controlling for and identifying various asset return regimes. Asset returns respond uniformly to fiscal policy shocks since the 1900's.

Keywords: TVP-VAR, countercyclical fiscal policy, stock prices, house prices

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1. Introduction

Fiscal policy shocks exert wide reaching effects, which include movements in asset returns. The policy response to the 2008/09 Great Recession generated enormous government bailouts of "too big to fail" businesses. These bailouts produced an expansion of the US government budget deficit, requiring either larger taxes or higher sovereign debt. The 2008/09 Great Recession captures one outlier with respect to the interaction between asset returns and fiscal policy. These outliers, together with a changing fiscal stance over the business cycle, highlight the nonlinear relationship between asset returns and fiscal policy. We contribute to the existing literature by analysing the simultaneous effect of fiscal policy shocks on asset returns as well as of asset return shocks on fiscal policy. We depart from the existing literature that uses vector autoregressive (VAR) models that average out period effects. Agnello et al. (2013) provide the exception, using a time-varying-transition-probability (TVTP) Markov-switching (MS) framework to adjust US fiscal policy for asset prices.

This study -- the first, as far as we know, analyses the interaction of fiscal policy and asset returns for the US in a parsimoniously restricted time-varying multivariate model, thus controlling for nonlinearities. The time-varying results capture key nonlinear effects such as explosive asset prices and various regimes switches. We use the generalised sup augmented Dickey-Fuller (GSADF) test (Phillips et al., 2015) to identify periods of explosive behaviour in asset returns. We also implement a Markov-switching framework to identify multiple asset return regimes (bull vs. bear markets). Both tests justify the use of our nonlinear VAR framework. We use a long history of annual data from 1890 to 2013 to capture important events and regimes in the movement in asset returns and control for changes in the fiscal stance. The results of the TVP-VAR show that both shocks to house and equity returns exert a larger effect on fiscal policy over time. Asset return shocks during the early 1900's did not generate much of an effect on fiscal policy. This outcome changed significantly since the 1970's. The response of asset returns to fiscal policy shocks depend on the duration of the shock. House returns react negatively to an increase in the fiscal primary surplus to GDP in all periods after the fiscal shock. Equity returns decrease in the initial periods after the fiscal shock. The equity return response to a fiscal policy shock, however, dissipates after one year.

Agnello et al. (2012) provide a brief summary of why the interaction between fiscal policy and asset prices are important. Two main channels exist through which asset prices affect fiscal policy: a direct channel where tax revenue increases in line with an increase in asset prices and an indirect channel where an increase in asset prices boost consumer confidence and, hence, increases aggregate demand. Eschenbach and Schuknecht (2002) show that asset prices affect fiscal balances mainly through the revenue channel; capital gains, turnover related taxes, and wealth effects and their effect on consumption affect the fiscal balance. That study, for 17 OECD countries, shows that, on average, a 10% change in real estate and stock prices exert a similar effect on the fiscal balance as a 1% change in output.

Liu, Mattina, and Poghosyan (2015) argue that the calculation of structural fiscal balances not only requires adjustment for the business cycle but also adjustment for asset price cycles, which do not fully synchronize with the business cycle. Thus, failure to account for asset price cycles in the calculation of the structural fiscal balance leads to a misleading signal about the underlying fiscal stance of the government. Further, asset price increases cause fiscal revenue to increase temporarily, which causes spending to respond to the higher revenue and produces a pro-cyclical policy stance.[1] Tagkatakis (2011b) re-enforces this

[1] Jaeger and Schuknecht (2004) and Morris and Schuknecht (2007) also support the view that the asset price cycle affects the structural fiscal balance differently than the business cycle.

revenue channel, finding that higher asset prices positively affects the primary balance through both higher revenue and lower spending.[2]

Tavares and Valkanov (2001) show that US fiscal policy affects asset price volatility in a magnitude at least as large as the effect of monetary policy in a VAR model. An increase in tax receipts reduces expected stock returns significantly while government spending does not significantly affect expected stock returns.

The increase in revenue due to an increase in asset prices may create a superficially higher revenue stream (possibly procyclical). This cyclically higher revenue may create the illusion of permanently higher revenue and as a result, may lead to higher fiscal spending. Budget balance calculations should, thus, strip the effects of asset prices on revenue to give an indication of structural revenue. Agnello et al. (2013) do this exactly in a TVTP-MS framework. They show that taxes adjust nonlinearly to asset prices, but spending remains neutral during an asset price cycle. This finding provides the most important justification for our contribution in a simultaneous systems model. We believe that the interaction between government taxation and spending proves important and, thus, requires study with movements in asset prices.

Furthermore, Agnello et al. (2012) use a variety of methods (linear and nonlinear) to study the response of fiscal policy to asset prices in the US and the UK. They show that asset prices do not affect primary spending. Taxes, however, do experience a significant effect - taxes decrease when stock prices rise and increase when house prices increase. Their nonlinear model shows that during economic downswings, fiscal policy expands and offsets some of the reduction in wealth typical during recessions, while a fall in wealth associates with a tax cut.

[2] Tagkalakis (2011a) reports that higher asset price volatility increases the volatility of government revenue, exacerbating the problem identified by Liu, Mattina, and Poghosyan (2015).

Apart from government's reaction to asset prices, Afonso and Sousa (2011) show that unexpected changes in fiscal policy increase the variability in asset prices. Aye et al. (2012) also study the effect of unexpected changes in fiscal policy on asset prices for South Africa. They show that fiscal policy announcements reduce the effects of fiscal policy shocks on asset prices in a sign restricted VAR.

Agnello and Sousa (2013) analyse the effect of fiscal policy in a panel VAR of ten industrialised countries. They show that a fiscal expansion exerts a negative effect on stock and house prices for the US.

Finally, our approach is similar to Gupta et al. (2014). They estimate a TVP-VAR for South Africa and show that a nonlinear relationship exists between fiscal policy and asset prices. They identify two significant regimes where this relationship changed. They demonstrate that fiscal expansions reduced both asset and house prices from the 1970's until the mid-1990's. Since 2000, however, asset prices increased in response to fiscal expansions. One explanation argues that fiscal policy during the 2000-2010 period was conducted in a sustainable and countercyclical way prompting consumers and firms to trust that fiscal policy will be used to stimulate aggregate demand during recessions and to save during economic expansions. Increasing taxes could limit the consolidation process, especially when it reduces asset prices (see Aye et al., 2012). As expected, an increase in asset prices reduced deficits.

2. Empirical methodology

To motivate the TVP-VAR model, we test for nonlinearities in asset price movements. We use two methods for this purpose: A generalised sup ADF test and a Markov-switching regression. The GSADF tests for bubbles and identifies the time origin of bubbles (Phillips et al., 2015). The test can detect bubbles in long time-series data. The GSADF test is a recursive right-tailed ADF test with flexible windows, where both the start and end points change.[3] We

[3] Phillips et al. (2014) describes the technical details of the test.

also use MS-AR regressions to complement the GSADF test. The MS-AR regressions identify the various bull and bear regimes.

In sum, we use the MS-AR models and the bubbles tests to identify bear and bull markets and possible periods of irrational exuberance in the two asset markets, independent of any other variables. Then we relate the time-varying impulse responses of the TVP-VAR model over these periods to identify any specific patterns in the impulse responses over the identified periods of bear or bull markets and bubbles. That is, do the impulse responses for a shock to budget deficit as a percentage of GDP (asset prices) on the asset prices (budget deficit as a percentage of GDP) differ under bear or bull markets and during bubbles.

Once we verify the presence of nonlinearities and identify periods of asset return exuberance, we use a TVP-VAR to analyse shocks over time and over the identified regimes. Time-varying impulses allow us to study the evolution of fiscal policy shocks to asset returns and vice versa. Time variation comes from both the parameters and the variance covariance matrix of the model's innovations. This reflects simultaneous relations among variables of the model and heteroscedasticity of the innovations (Primiceri, 2005). To accomplish this, we use a Monte Carlo Markov Chain (MCMC) algorithm to estimate the coefficients and the multivariate stochastic volatility. Researchers commonly estimate time variation (see Sims 1993, Stock and Watson 1996, and Cogley and Sargent 2001). These studies, however, impose restrictions on the variance covariance matrix that should evolve over time. Most of these models limit themselves to reduced-form models that can only describe and forecast data (Primiceri, 2005). With drifting coefficients, one essentially captures the learning process and possible nonlinearities or time variation in the lag structure of the model. The multivariate stochastic volatility can capture possible heteroscedasticity of the shocks and nonlinearities in the simultaneous relations among the variables of the model.

Following Nakajima (2011), this paper estimates a time-varying parameter VAR model with stochastic volatility of the form:

$$y_t = c_t + B_{1t} y_{t-1} + \ldots + B_{st} y_{t-s} + e_t, \quad e_t \sim N\left(0, \Omega_t\right), \tag{1}$$

where $t = s+1, \ldots, n$, y_t is a $(k \times 1)$ vector of observed variables, $B_{1t}, \ldots, B_{st}$ are $(k \times k)$ matrices of time-varying coefficients, and Ω_t is a $(k \times k)$ time-varying covariance matrix. We assume a recursive identification scheme by the decomposition of $\Omega_t = A_t^{-1}\Sigma_t\Sigma_t A_t^{'-1}$, where A_t is a lower-triangle matrix with diagonal elements equal to one, and $\Sigma_t = \mathrm{diag}\left(\sigma_{1t}, \ldots, \sigma_{\kappa t}\right)$. Let us define β_t as the stacked row vector of $B_{1t}, \ldots, B_{st}$; a_t is the stacked row vector of the free lower-triangular elements of A_t; and $h_t = (h_{1t}, \ldots, h_{kt})$ where $h_{jt} = \log \sigma_{jt}^2$. We assume that the time-varying parameters follow a random-walk process. Thus,

$$\begin{array}{l} \beta_{t+1} = \beta_t + \upsilon_{\beta t}, \\ a_{t+1} = a_t + \upsilon_{at}, \\ h_{t+1} = h_t + \upsilon_{ht}, \end{array} \qquad \begin{pmatrix} \varepsilon_t \\ \upsilon_{\beta t} \\ \upsilon_{at} \\ \upsilon_{ht} \end{pmatrix} \sim N\left(0, \begin{pmatrix} I & 0 & 0 & 0 \\ 0 & \Sigma_\beta & 0 & 0 \\ 0 & 0 & \Sigma_a & 0 \\ 0 & 0 & 0 & \Sigma_h \end{pmatrix}\right), \tag{2}$$

where $t = s+1, \ldots, n$, $e_t = A_t^{-1}\Sigma_t\varepsilon_t$, Σ_a and Σ_h are diagonal, $\beta_{s+1} \sim N\left(\mu_{\beta o}, \Sigma_{\beta o}\right)$, $a_{s+1} \sim N\left(\mu_{ao}, \Sigma_{ao}\right)$, and $h_{s+1} \sim N\left(\mu_{ho}, \Sigma_{ho}\right)$.[4] We use Bayesian inference to estimate the TVP-VAR model via MCMC methods. The MCMC methods assess the joint posterior distributions of the parameters of interest under certain prior probability densities. We assume the following priors, as in Nakajima (2011): $\Sigma_\beta \sim IW(25, 0.01I)$, $(\Sigma_\alpha)_i^{-2} \sim G(4, 0.02)$, $(\Sigma_h)_i^{-2} \sim G(4, 0.02)$, where $(\Sigma_\alpha)_i^{-2}$ and $(\Sigma_h)_i^{-2}$ are the i-th diagonal elements in Σ_α and Σ_h, respectively, and IW and G denote the inverse Wishart and the

[4] For a comprehensive analysis of the TVP-VAR methodology and the estimation algorithm, refer to Nakajima (2011).

gamma distributions, respectively. For the initial set of the time-varying parameter, we set flat priors such that: $\mu_{\beta o} = \mu_{ao} = \mu_{bo} = 0$ and $\Sigma_{\beta o} = \Sigma_{ao} = \Sigma_{bo} = 10 \times I$.

The VAR exhibits a recursive identification scheme, where depending on the ordering of the variables, contemporaneous shocks are zero. We order the variables from most exogenous to least exogenous with house returns first, followed by the primary surplus, and then equity returns.[5] This ordering is theoretically justified (Gupta et al., 2014), since it implies that the house price does not respond contemporaneously to fiscal policy and equity price shocks, while fiscal policy reacts with a lag to equity price shocks. Thus, the equity return appears third in the ordering after the house return and measure of fiscal policy.

3. Data description

Our variables include the real US house price index (RHP), real Standard and Poor's S&P500 (RSP) index and the US primary surplus as a percent of Gross Domestic Product (BB).[6] We use annual data from 1890 to 2013, because data only occur at this frequency over this long-sample. We use yearly growth rates (log-differences) for the real house and stock returns (RHR and RSR respectively), which generates a total of 123 observations, covering the period of 1891-2013. We plot the RHP and RSP series in Figure A1 in the Appendix. Both series exhibit non-stationary behaviour at the conventional 5-percent significance level. The level of BB and the transformation of the two asset prices into returns generate stationary series based on standard unit-root tests (i.e., Augmented-Dickey-Fuller (ADF) (Dickey and Fuller, 1981), Phillips-Perron (PP) (Phillips and Perron, 1988), Dickey-Fuller with Generalised-Least-Squares-detrending (DF-GLS) (Elliott et al., 1996), and the Ng-Perron modified version of the PP (Ng and Perron, 2001)), which we report in Table A1 in the

[5] We use two lags for the VAR as given by the Schwarz information criterion (SIC). We interchange the variable ordering to evaluate the robustness of the results. The results mainly remain the same. Results are available from the authors on requested.
[6] The RHP and RSP data come from the Online Data section of Robert Shiller's website: http://www.econ.yale.edu/~shiller/data.htm while the nominal surplus and GDP data come from the Global Financial Database.

Appendix. We standardise all variables to ensure that we can easily compare the results across the two asset returns. Summary statistics of the non-standardized variables, however, appear in Table A2 in the Appendix. All variables exhibit non-normal behaviour, providing initial motivation to include stochastic volatility in the model.

Our univariate MS regressions identify two regimes for both house and equity returns: A high volatility state (regime 1), which is a bear market, and a low volatility state (regime 2), which is a bull market (see Simo-Kenge et al., forthcoming and references therein). Table 1 summarises the results of the MS-AR regressions. The observed volatility is higher for both house and equity returns during bear markets compared to bull markets. The expected duration in the bear market for house returns is about 80 years compared to about 34 years during the bull market.[7] The expected duration is almost equal for both states for equity returns. For stock returns, the duration for bear and bull markets proves much shorter with 2 and 2.5 years, respectively

Figure 1 plots the regime probabilities of being in a bull market for both stock and house returns. Stock and house return bull markets nearly coincide with each other prior to 1950. That does not occur in the post-1950 period. The housing market experienced a long low volatility period from 1950 until the collapse of the housing market leading up to the financial crisis in 2008.

Similar to Pavlidis et al. (2013), the GSADF test for the real house price shows signs of explosive behaviour during the early 2000's until around 2007 (see Figure 2). The correction from this bubble helps to explain the 2008/09 Great Recession. A real stock price bubble (see Figure 3) emerges from 1995 to 2003. Phillips et al. (2015) use a higher

[7] We also estimated a three-regime model for housing returns. The estimation broke down, however, due to a singular covariance, which, in turn, implied non-unique coefficients. This outcome possibly reflects an over parametrized model with just 123 observations. But, in general, the estimations identified two bear regimes, which is not surprising. If one examines the real house price series in Figure A1, it remains relatively flat until 1997 with a sharp decline in 1916, which resulted in low real house prices until 1940. Complete details of these results are available upon request from the authors.

frequency series for the S&P data over a similar time period and identifies a bubble from 1995 to 2008. Both asset prices in real terms reveal an explosive and unsustainable bubble that preceded the Great Recessions.

4. Results

Having identified key periods of asset return exuberance and potential bubbles, we turn to the TVP-VAR model to analyse the response of asset returns to fiscal policy as well as the response of fiscal policy to asset return shocks.[8]

We generate the posterior estimates after drawing 10,000 samples, with the first 1,000 draws discarded. Table 2 reports these posterior estimates for the means, along with those for the standard deviations, the 95 percent credibility intervals[9], the convergence diagnostic (CD) due to Geweke (1992), and the inefficiency factors.[10] The 95 percent credibility intervals include the estimates for the posterior means, and the CD statistics do not allow us to reject a null hypothesis of convergence to the posterior distribution at the 5-percent significance level. Furthermore, the inefficiency factors are relatively low. We can thus conclude that the MCMC algorithm efficiently produces the posterior draws. This, in turn, also implies that even with annual data (instead of the traditional quarterly data used in this literature), our model does not suffer from imprecision in the parameter estimates and the generated impulse responses (which we discuss below).

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[8] Based on the suggestions of an anonymous referee, to motivate our decision to conduct the impact of asset returns shocks on fiscal policy, besides them being affected by the fiscal policy shock, we conducted Granger causality analysis for the full-sample, as well as, sub-samples identified by Bai and Perron's (2003) multiple structural break tests. Our results, which are available upon request from the authors, provided ample evidence of fiscal policy not only causing asset returns, but also being affected in turn, by real stock and housing returns.

[9] We use credibility intervals in the Bayesian paradigm as opposed to 'confidence' intervals, which belong in the frequentist realm.

[10] Geweke (1992) suggests the comparison between the first n_0 draws and the last n_1 draws, dropping out the middle draws, to check for convergence in the Markov chain. We compute the CD statistics as follows: $CD = (\bar{x}_0 - \bar{x}_1)/\sqrt{\frac{\hat{\sigma}_0^2}{n_0} + \frac{\hat{\sigma}_1^2}{n_1}}$, where $\bar{x}_j = (\frac{1}{n_j}) \sum_{i=mj}^{mj+nj-1} x^{(i)}$, $\bar{x}_j$ is the i-th draw, and $\frac{\hat{\sigma}_j^2}{n_j}$ is the standard error of $\bar{x}_j$, respectively, for j=0, 1. If the sequence of the MCMC sampling is stationary, it converges to a standard normal distribution. We set $m_0 = 1, n_0 = 1000, n_1 = 5001$ and $n_2 = 5000$. $\hat{\sigma}_j^2$ is computed using a Parzen window with bandwidth $(B_m) = 500$. The inefficiency parameter is defined as $1 + 2\sum_{i=1}^{Bm} \rho_i$, where ρ_i is the sample autocorrelation at lag s, which is computed to measure how well the MCMC chain mixes. Figure A1 in the appendix gives the convergence output with moments and posterior distributions.

Figure 4 plots the posterior estimates of stochastic volatility for each of the variables used in the TVP-VAR. The top row plots the actual time series while the bottom row plots the posterior estimates of the stochastic volatility of each series. The stochastic volatility of equity returns remain more or less constant throughout our sample and remain high. The most prominent point in the posterior estimate of the primary surplus stochastic volatility occurs during WWII (1939-1950), which matches what we observe in the evolution of fiscal policy during that period. Smaller movements in volatility also occur for WWI and even smaller, but noticeable, movements occur during the Great Recession. Stochastic volatility for house returns achieved its highest level at the end of the 19th century and increased again from 2000 to 2007. The stochastic volatility between fiscal policy and asset returns do correspond over specific periods. In contrast, Simo-Kenge et al. (forthcoming) show that monetary policy and asset return volatility correlate over specific periods. The higher volatility in house returns during the beginning of the 1900's to around 1940 corresponded to relative low volatility in interest rates. The non-constant posterior estimates for stochastic volatility justify our use of the model with stochastic volatility as opposed to one with constant volatility in the structural shocks.

The impulse responses of Figure 5[11] show how the primary balance reacts to asset return shocks over time. We have nine impulse horizons from 1892 to 2013. We summarise, however, only impulses at years 1, 3, 6, and 9. Both house and equity return shocks lead to higher primary surpluses over time. This matches the constant parameter VAR of Figure A3 for the first couple of years (see Appendix). We expect increases in house and equity returns, a priori, to increase automatically the revenue collected from these tax bases. At the same time, we also expect countercyclical spending from government during periods of buoyant growth. Interestingly, the effects of asset return shocks on the primary balance exhibit

[11] See the Appendix for impulse response with 68-percent credibility intervals.

stability since the 1990's. This indicates that the government did not react differently to asset return shocks during two significant periods of asset price movements: the dot com bubble in 2000 and the 2008/09 Great Recession. The primary balance reacts more to stock return shocks contemporaneously compared to house return shocks, indicating that movements in equity markets provide either a stronger source of revenue or a signal to government of changes to the economic cycle. Of course, tax revenue changes occur only for assets sales, realized gains or losses. Since houses contain both consumption and investment components, whereas equities include only the investment component, realized gains or losses will occur more frequently for equities than for houses. Nonetheless, the effect of stock return shocks on the primary balance quickly die out after one period.

Both house and equity returns respond negatively to a fiscal surplus shock contemporaneously, a finding similar to one in Gupta et al., (2014). The response of house returns to a fiscal shock lasts up to three years whereas the response of stock returns to a fiscal policy shock lasts only one year. This is similar to Figure A4 (see Appendix) for a constant parameter VAR, although the real stock return increases initially for a period before turning negative as well, opposite to what the TVP-VAR model generates. This negative response to fiscal shocks reflects a standard market response. That is, when the primary surplus increases, or more likely the primary deficit decreases, the supply of government debt falls, driving up the price of government debt and lowering its interest rate. As such, the quantity demanded of government debt declines, moving to the housing and equity markets, lowering the returns in these markets. Interestingly, we see little volatility in the impulse responses over time. These responses increase smoothly, although at differing rates, over time (see Figures 5 and 6). The exception, the real stock return response to fiscal shocks remains relatively uniform since the 1900's, albeit at different levels (see Figure 6).[12]

Finally, we performed the MS and bubbles tests on the univariate data to see if the behaviour of asset prices and fiscal policy differed somewhat between bear and bull markets as well as during bubbles in the two asset markets. The smoothness of the impulses in Figures 5 and 6 suggest that the impulses do not differ dramatically between bear and bull markets and/or during bubbles.

5. Conclusion

This paper studies the dynamic interaction between fiscal policy and asset returns in a TVP-VAR setup. This method addresses potential nonlinearities between fiscal policy and asset returns and controls for exuberant periods of asset returns as well as a changing fiscal policy stance. The use of the TVP-VAR is motivated by two tests - a GSADF tests that detects and dates bubbles and a Markov-switching regression that identifies multiple asset return regimes.

After controlling for time-varying stochastic volatility, the results show that positive asset return shocks increase the primary surplus. The response of fiscal policy to asset returns increases every year from the 1890's into the early 1990's. Fiscal policy's reaction to asset return movements becomes fairly stable after 1990. The findings also show that fiscal policy did not overreact to drastic changes in asset returns during the 2008/09 financial crisis.

[12] Based on the recommendations of the two anonymous referees, we conducted the following robustness tests: (a) Instead of the primary surplus to GDP ratio, we used a four variables VAR, where we included real GDP growth along with the real budget surplus; (2) In the benchmark model of the three variables, we added real interest rate (with the nominal interest rate and CPI used to compute the inflation rate, obtained from the data segment of Professor Robert J. Shiller's website); and, (3) We conducted the analysis over the period of 1929-2013 (due to data availability, with the same obtained from the US Census Bureau) by using tax-GDP and government expenditure-GDP ratios instead of the budget surplus to GDP ratio. Our results were, however, qualitatively similar to those reported in the paper. Under (3), when we analysed the period of 1929-2013, we found that tax-GDP ratio decreased asset returns more strongly than the increase in asset prices resulting from an increases in the government expenditure-GDP ratio. In addition, increase in asset returns also increased tax-GDP ratio relatively more than the reduction in government expenditure-GDP ratio. This explains why we see a fall in asset returns following an increase in budget surplus to GDP ratio, and an increase in the budget surplus to GDP ratio due to increases in asset returns, when the analysis was conducted in the three variables VAR over the same period of 1929-2013. Given that including all these results would lead to an unprecedented increase in the length of the paper, we decided to exclude them for the sake of brevity. However, complete details of these results are available upon request from the authors.

The results also show that asset returns react negatively to an increase in the primary surplus. Stock returns, however, increase in response to a fiscal shock after a year. The TVP impulse responses remain fairly constant over time. This suggests a constant reaction of asset returns to fiscal policy shocks since 1900. Although house returns decrease due to an expansion in fiscal policy, stock returns only decrease during the first year after the shock.

An issue that needs to be highlighted at this stage is anticipation effects and hence, nonfundamentalness. Nonfundamentalness arises in situations in which the information set of the econometrician is smaller than that of the economic agents (Lippi and Reichlin, 1994). Given this, Ramey (2011) emphasizes that neglecting anticipation effects in fiscal VARs can cause the impulse responses to be biased, and hence, suggests the need to include news about future fiscal policy to overcome this problem, as done in Berg (2015) for instance using professional forecasts. However, we are unable to accommodate for nonfundamentalness in our study, as professional forecasts on the fiscal policy variables and asset prices are clearly not available over the historical period of 1891-2013. This can be considered as the trade-off involved in looking at the historical evolution, instead of recent periods only, of the relationship between fiscal policy and asset returns.

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Table 1: MS-AR estimates

	Real house returns (RHR) (MS-AR(3))	Real stock returns (RSR) (MS-AR(2))
Regime dependent intercept		
$\boldsymbol{\mu_1}$	-0.174	-7.204
$\boldsymbol{\mu_2}$	0.388	10.710***
Standard errors		
$\boldsymbol{\sigma_1}$	2.143***	2.855***
$\boldsymbol{\sigma_2}$	0.710***	2.423***
Regime 1: Parameter estimates		
$\boldsymbol{\beta_{1,1}}$	0.008	0.275
$\boldsymbol{\beta_{1,2}}$	0.018	-0.431**
$\boldsymbol{\beta_{1,3}}$	-0.168	
Regime 2: Parameter estimates		
$\boldsymbol{\beta_{2,1}}$	1.027***	-0.189
$\boldsymbol{\beta_{2,2}}$	-0.171	0.035
$\boldsymbol{\beta_{2,3}}$	-0.165	
Fit		
Log-likelihood	-364.256	-514.627
Transition probabilities		
Bear	0.988	0.549
Bull	0.971	0.606
Expected duration		
Bear	80.328	2.219
Bull	34.806	2.541

Notes:*,**,*** indicates significance at 10%, 5% and 1% respectively

Table 2: Selected estimation results

Parameter	Mean	Stdev	95% Intervals	CD	Inef
$\left(\sum \boldsymbol{\beta}\right)_1$	0.0023	0.0003	[0.0018,0.0029]	0.018	5.68
$\left(\sum \boldsymbol{\beta}\right)_2$	0.0023	0.0003	[0.0018,0.0029]	0.074	7.76
$\left(\sum \boldsymbol{\alpha}\right)_1$	0.0056	0.0016	[0.0034,0.0095]	0.899	34.23
$\left(\sum \boldsymbol{\alpha}\right)_2$	0.0054	0.0014	[0.0033,0.0088]	0.628	23.73
$\left(\sum \boldsymbol{h}\right)_1$	0.3539	0.0960	[0.1950,0.5580]	0.089	35.15
$\left(\sum \boldsymbol{h}\right)_2$	0.8358	0.1542	[0.5611,1.1567]	0.164	14.82

Figure 1: MS-AR bull market probabilities for real stock and house returns

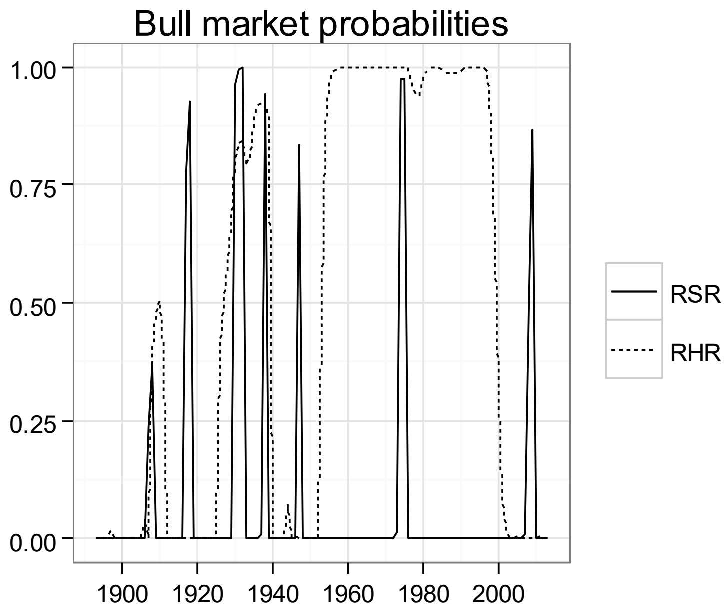


Figure 2: GSADF test for house returns

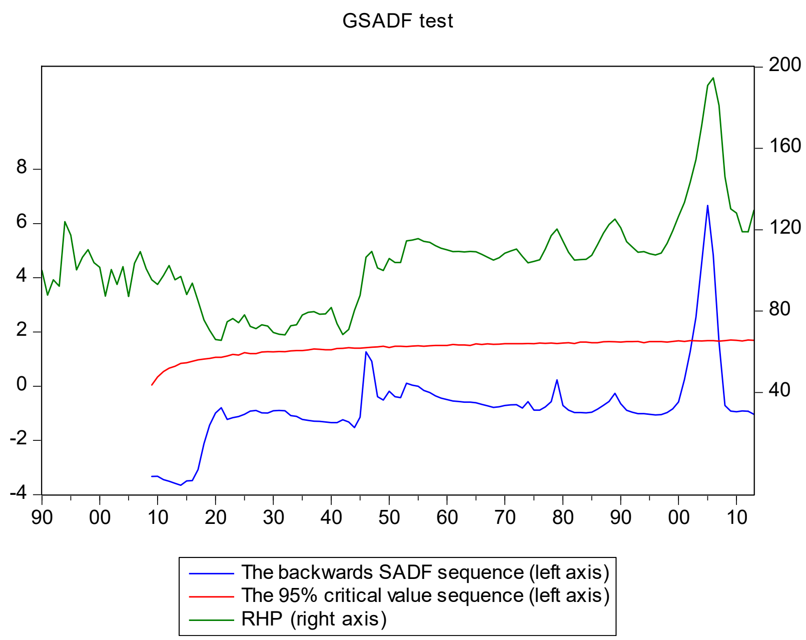


Figure 3: GSADF test for equity returns

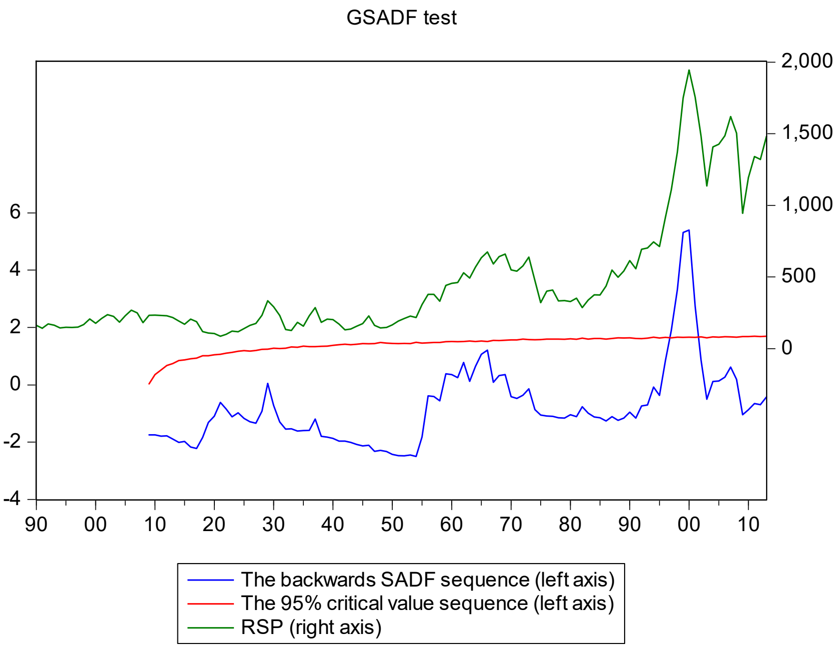


Figure 4: Posterior estimates for the stochastic volatility of the structural shock

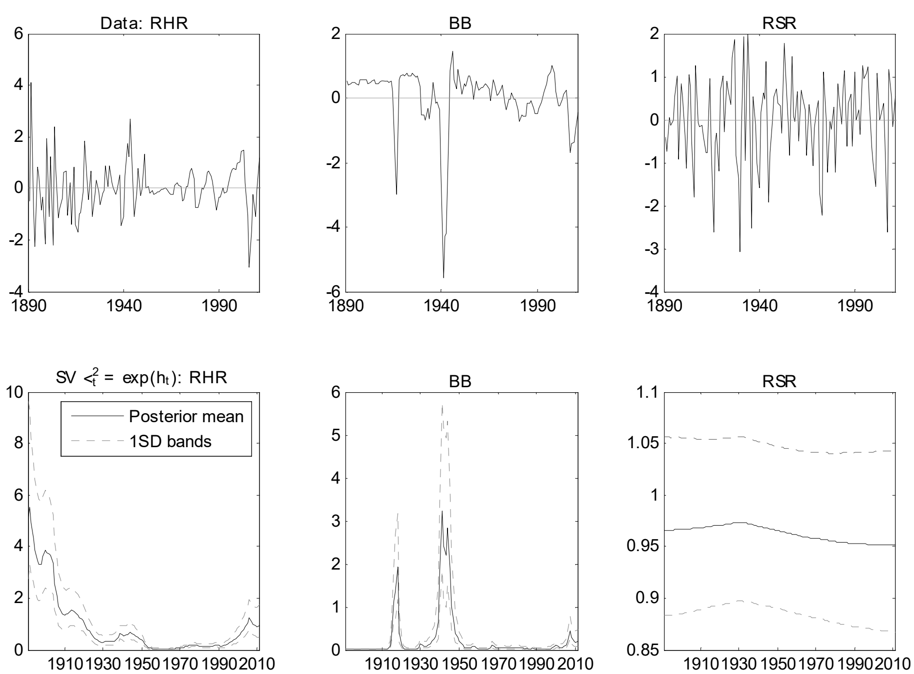


Figure 5: Median response the primary balance to an asset return shocks

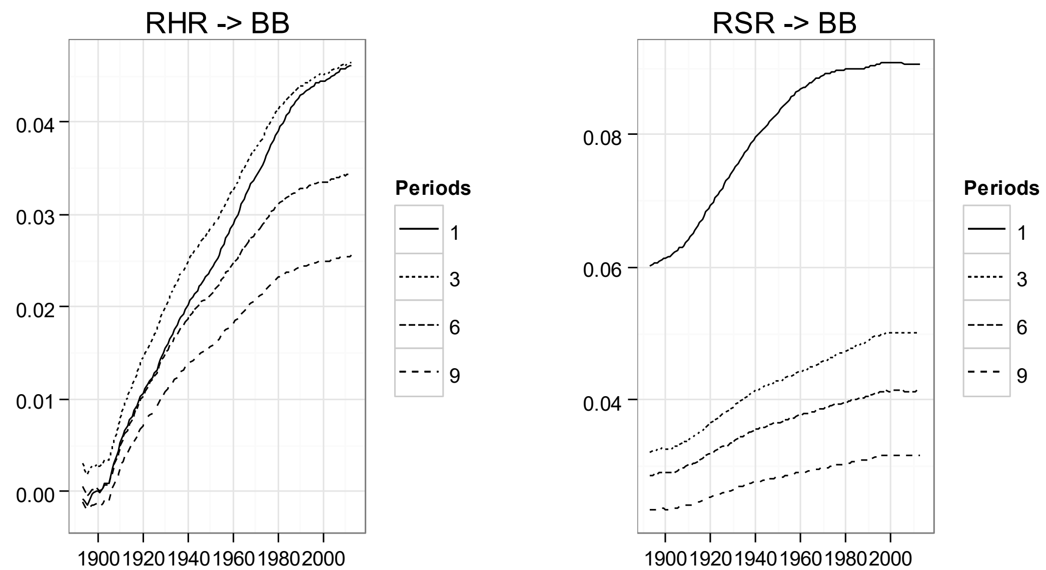
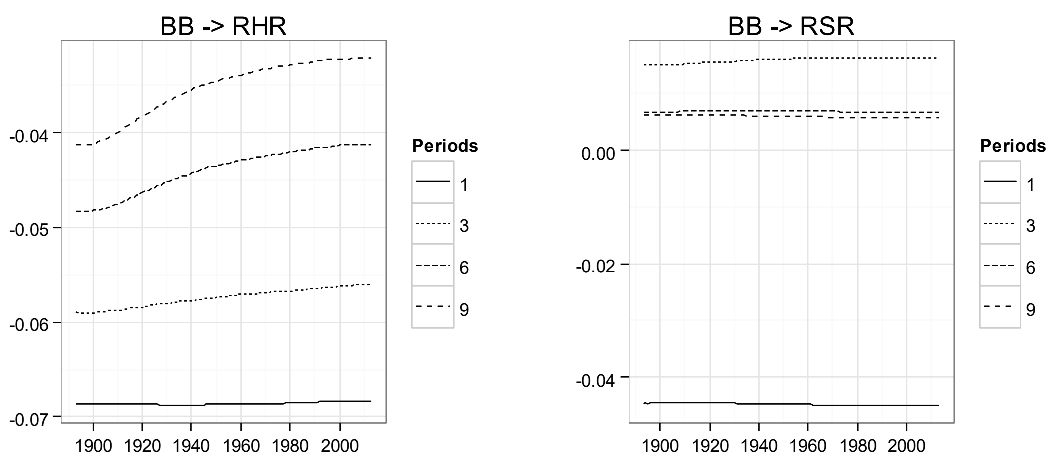
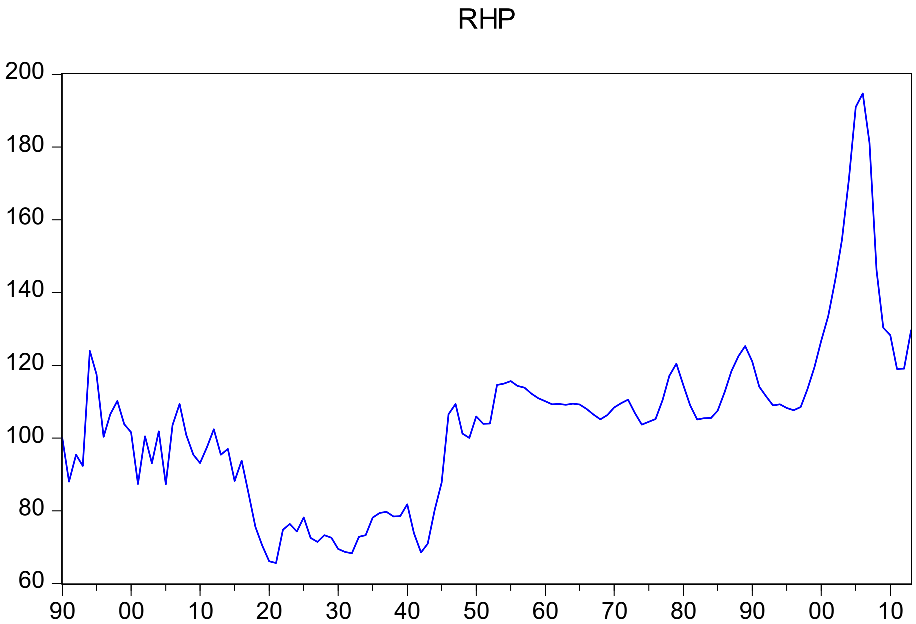


Figure 6: Median response of asset returns to a primary surplus shock



Appendix:

Figure A1: Plots of Real House Price (RHP) and Real Stock Price (RSP)



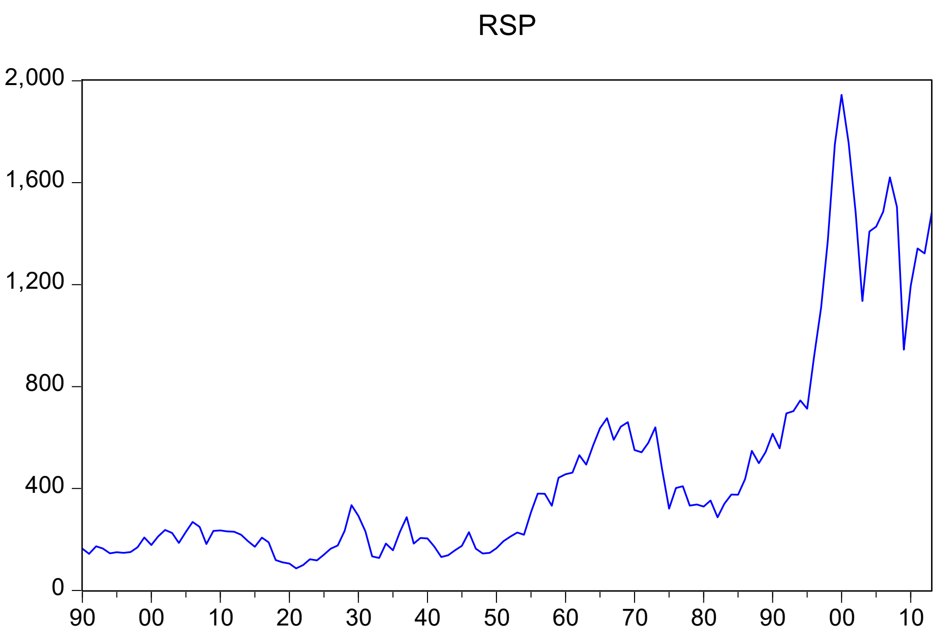
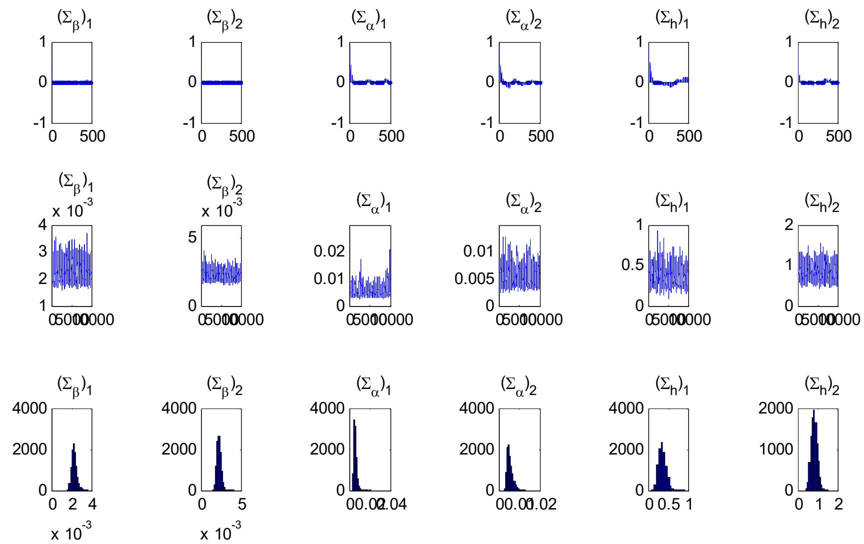
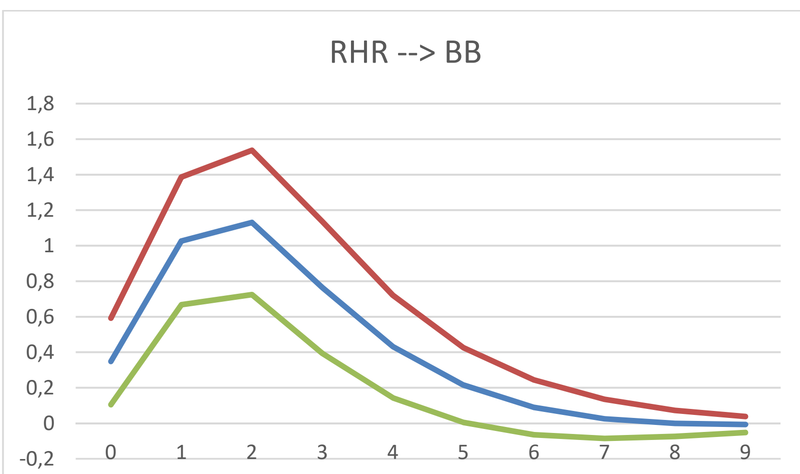


Figure A2: Moments and posterior distribution



Notes: Sample autocorrelation (top chart), sample paths (middle chart) and posterior densities (bottom chart)

Figure A3: Constant VAR with 68% confidence interval: Asset return shocks



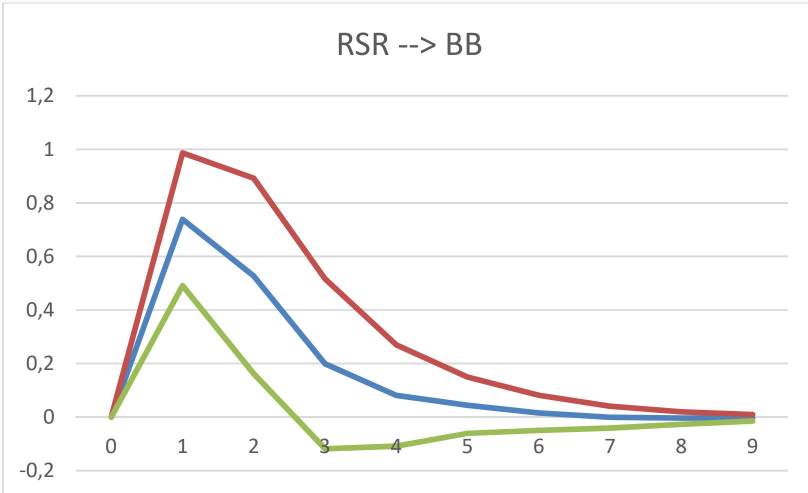
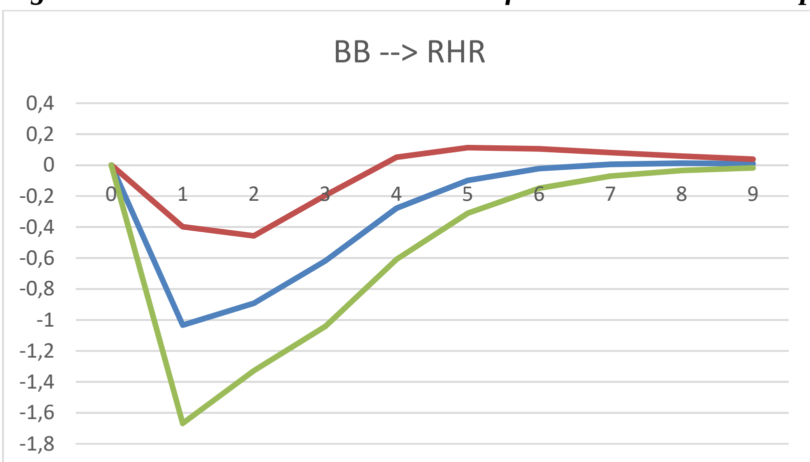


Figure A4: Constant VAR with 68% confidence interval: Fiscal policy shocks



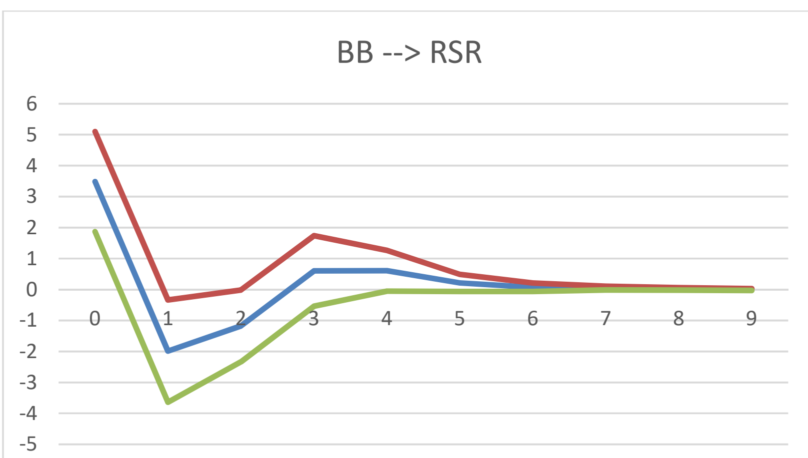


Figure A5: Model response with credibility intervals for 1 year horizon

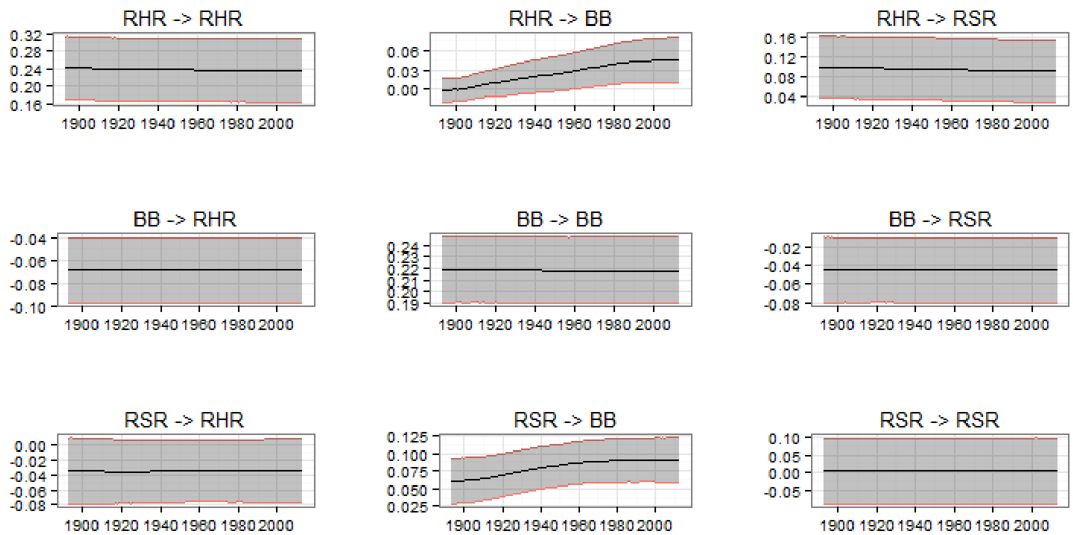


Figure A6: Model response with credibility intervals for 3 year horizon

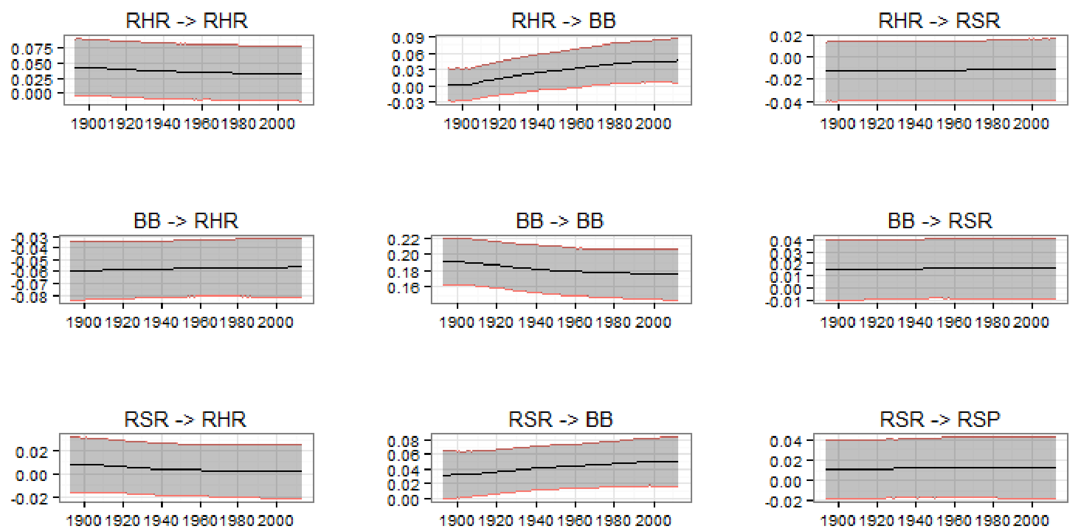


Figure A7: Model response with credibility intervals for 6 year horizon

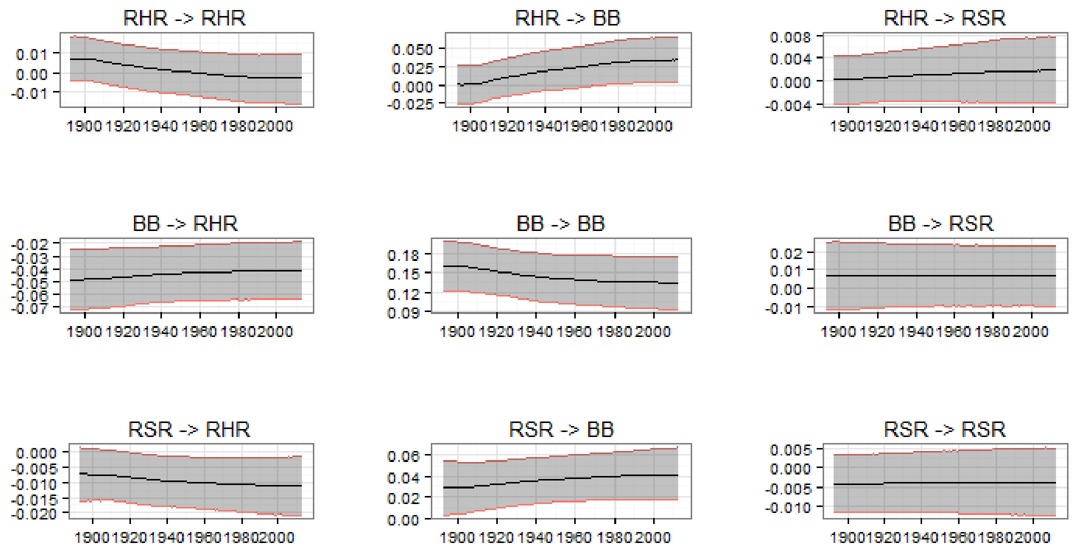


Figure A8: Model response with credibility intervals for 9 year horizon

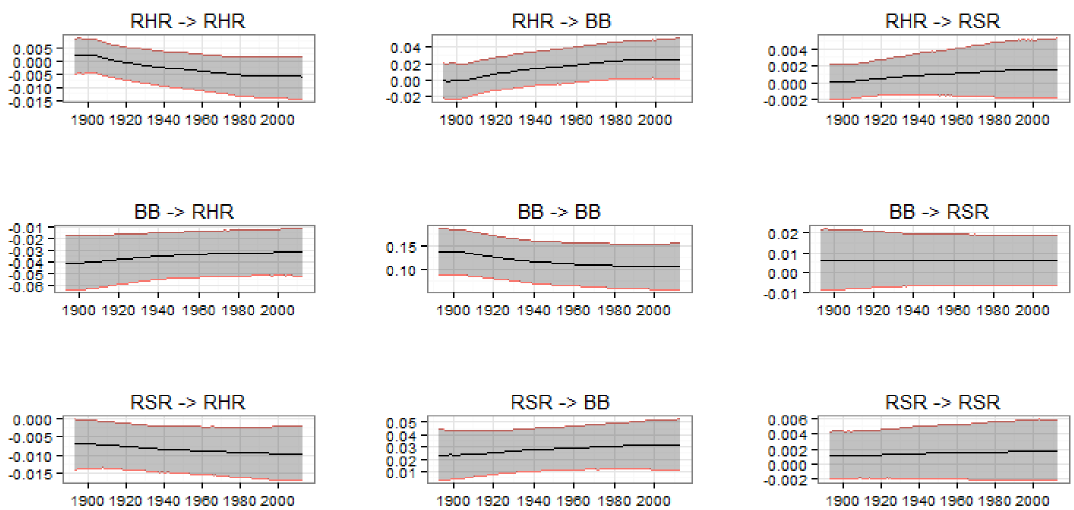


Table A1: Unit Root Results

	Constant				Constant + Trend			
	Level		First Difference		Level		First Difference	
BB								
ADF	-5.3720	***	-		-5.4470	***	-	
DF-GLS	-4.9392	***	-		-5.4533	***	-	
PP	-3.7173	***	-		-3.7462	***	-	
Ng-Perron	-45.4748	***	-		-53.9040	***	-	
RHP								
ADF	-1.8050		-10.3847	***	-2.5088		-10.3449	***
DF-GLS	-1.8170	*	-3.0896	***	-2.2549		-4.9458	***
PP	-1.7932		-10.3902	***	-2.7006		-10.3501	***
Ng-Perron	-6.8997	*	-15.7581	***	-10.0853		-31.6338	***
RSP								
ADF	-0.7520		-10.6227	***	-2.4978		-10.6041	***
DF-GLS	-0.1526		-8.5429	***	-2.2483		-10.0368	***
PP	-0.6995		-10.6536	***	-2.5801		-10.6373	***
Ng-Perron	-0.3029		-57.3289	***	-9.6890		-60.4707	***

Note: *** and * indicates rejection of the null of unit root at 1 percent and 10 percent levels of significance respectively.

Table A2: Summary Statistics

	BB	RHR	RSR
Mean	-2.2899	0.2108	1.7949
Median	-0.7508	0.0728	2.6222
Maximum	4.2879	29.5108	38.1463
Minimum	-27.4693	- 21.4550	- 54.9023
Std. Dev.	4.5299	7.0940	18.3697
Skewness	-3.0551	0.3749	-0.6145
Kurtosis	14.5459	5.3887	3.3993
Jarque-Bera	874.5433	32.1244	8.5569
Probability	0.0000	0.0000	0.0139
Observations	123	123	123

Note: Std. Dev: Standard deviation; Probability relates to the Jarque-Bera test which has a null of normality.