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ORIGINAL PAPER

The impact of the Brexit vote on UK financial markets: a synthetic control method approach

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Abstract
We estimate how the UK financial markets would have evolved if the Remain camp had won the referendum. To construct the counterfactual, we use the synthetic control method. Our results suggest that there would not have been any significant change in the development of the FTSE 100 Index in the medium to long term if there had not been a referendum. On the other hand, we find a significantly negative effect of 1.2 percentage points on the 10-year bond yield. Given the geopolitical circumstances in mid 2016, financial agents investing in the pound could have sought safer investment options represented by longer-term government bonds, which consequently could result in lower bond yields.

Keywords Brexit · Financial markets · Macroeconomic indicators · Synthetic control method

JEL Classification C10 · Q10 · Q18

1 Introduction

Economists and social scientists are often interested in the impact of events or interventions at an aggregate level on entities such as countries, regions or firms. The unexpected result of the Brexit vote on June 23, 2016 provides a unique opportunity to study the impact of this idiosyncratic event on selected financial indices. Moreover, the unpredictability of the vote allows us to use the synthetic control method (SCM) introduced by (Abadie and Gardeazabal 2003), which provides a data driven approach in which control units are systematically chosen as a weighted average of all relevant units that best fit the characteristics of the treatment unit. Using the

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weighted average approach precludes the extrapolation that is the typical basis of regression models (King and Zeng 2006).

There is limited literature related to the impact of the Brexit vote on the British economy. Born et al. (2019) use the SCM to evaluate the effect of the vote on the change in future expectations about the UK economy. They find that a large part of the cost of Brexit could be attributed to the gloomier average expectations of economic agents about the British economic future. Another contribution by Ramiah et al. (2017) using the event study methodology shows that the banking, travel and leisure sectors were negatively affected, with cumulative abnormal returns of—15.37% for the banking sector. Bloom et al. (2019) study the impact of Brexit on UK firms by using the major new survey. They stress out that the main impacts of Brexit are long-lasting increase in uncertainty, gradually reduced investment by about 11% and reduced UK productivity by about 2–5% over 3 years following the referendum. Breinlich et al. (2018) study stock market reactions to the referendum using a model of 'normal' stock returns which takes into account the differences in risk and other characteristics of stocks. Their result suggests that in the short run the initial drop in stock price was mainly driven by sterling depreciation and by fears of cyclical downturn. Another study related to the effect of the referendum on the stock market was conducted by Davies and Studnicka (2018). By using two-stage estimation process they examined how news of Brexit affected expectations in stock returns. They find that most firms had negative returns following news of the Brexit referendum. Consequently, they explore that there was a considerable heterogeneity in their changes relative to expectations. They contribute these heterogeneity to the firm's global value chain, with companies exposed to the EU and the UK performing worse. Our study indicates similar findings in the short term, however, given the monthly periodicity of our data we rather show the impact in medium to long term period.

The consequences of the Brexit vote on living standards was studied by Breinlich et al. (2017), who suggest that the referendum increased aggregate UK inflation by 1.7 percentage points within 1 year. Moreover, they find that the cost of inflation is evenly shared across the income distribution, but not across regions, with London the least affected and Scotland, Wales and Northern Ireland impacted the most. The effect of the Brexit vote on corporate loan origination was estimated by Berg et al. (2016). They propose a new matching strategy called Siamese twins matching, which is used to find appropriate counterfactuals for the UK market. The results show that the UK syndicated loan market dropped by 25% after the Brexit vote.

Deploying the SCM, we build a counterfactual world that shows how selected variables would have developed if the Brexit vote had not occurred. In this paper, we estimate the impact on the UK stock exchange, long-term government bonds and exchange rate. We approximate the aforementioned variables by long-term treasury yield (10-year gilt yield), Financial Times Stock Exchange 100 Index (FTSE 100 Index) and the real effective exchange rate (REER). We use monthly data from the OECD database to establish the donor pool of countries that best resembles the economic development of the UK before the vote. We contribute to the current literature on the economic impact of the Brexit vote by analysing the development of stock and bond markets and the real effective exchange rate.

As far as we know, this paper is first to analyse the impact of the Brexit vote on UK financial market using SCM. Our paper relates to the analysis of macroeconomic experiments at the aggregate level Billmeier and Nannicini (2013), Gathani et al. (2013) and Hosny (2012) and the literature on employing the synthetic control method. From the outset, we would like to stress that we are not testing a model of the UK economy. Instead, we attempt to establish a possible path of selected variables and the magnitude of the effect of the Brexit vote.

The results based on monthly data show that there would not have been any significant change in the development of the FTSE 100 Index if the referendum had not occurred. Since that companies listed on the FTSE 100 derive 76% of their revenue outside the UK, financial agents could take advantage of British currency depreciation and implement a wait-and-see approach (FTSE Russell 2017). As the Bank of England in their Quarterly Bulletin from 2016 Q3 points out (p. 146): ‘On 23 June UK-focused equity prices fell sharply. However since then, expectations for a broadly more accommodative stance of monetary policy from major central banks helped to reverse some of these moves. Equities performed strongly, credit spreads decreased, and volatility fell to record lows’ Manning (2016). This corresponds with the result for REER, which would not have declined in the case of no referendum. Finally, we find a significant effect on the 10 year bond yield, which would have been by 1.2 percentage points higher had there not been any referendum in long-term period. Similar magnitude of the the effect can be find after quantitative easing (QE) in February 2010 done by Bank of England. This event depressed 10-year bond yield by 100 basis points (Joyce et al. 2011).

In their study, Born et al. (2019) show that macroeconomic uncertainty explains close to half of the observed output loss in the UK after the Brexit vote. Adding the geopolitical circumstances in mid 2016, financial agents investing in GBP (British pound sterling) could seek safer options represented by longer term government bonds, which results in lower bond yields.[1] Moreover, as Manning (2016) reports in Quarterly Bulletin 2016 Q3 (p. 147) : ‘At longer maturities, contacts attributed some of the moves to changes in hedging activities of liability-driven investors such as insurers and pension funds.’

In general, the short term result corresponds to the phenomenon that the bond market plays its traditional role as a hedge for the equity market. This is mainly in line with empirical studies of Baur and Lucey (2009), Garcia and Tsafack (2011) and Ciner et al. (2013). Garcia and Tsafack (2011) study symmetric and asymmetric dependence between international equity and bond markets. They propose an alternative regime-switching copula model to estimate the dependency in the US, Canada, France and Germany stock and bonds markets. They find that dependence between international assets of the same type is strong in both regimes, especially in the asymmetric one, but weak between equities and bonds, even in the same country. This findings is in line with our results, which show that equity market stayed relatively unaffected compared to long-term government yield in the mid to long term period. The study conducted by Baur and Lucey (2009) analyses the existence

[1] The presidential election in the USA, the Syria crisis and the EU immigration policy discussion.

of flights from stocks to bonds and vice versa. They use daily continuously compounded MSCI stock and bond index returns of the US, the UK, Germany, France, Italy, Australia, Canada and Japan. Based on their empirical findings they show that flights are a common feature in crisis periods and often occur simultaneously across countries. Another study done by Ciner et al. (2013) investigate the return relations between equity, bonds, gold, currency and oil markets using data from both the US and the UK. Their primary result show that on average bond market plays its traditional role as hedge for the equity market. This is in line with previous study of Baur and Lucey (2009) and corresponds with our results in the short term.

In conclusion, the unexpected result of the Brexit vote helped to decrease long-term government yield and triggered a strong decline in REER. On the other hand, the equity market stayed relatively unaffected in the mid to long term thanks to the accommodative monetary policy of the Bank of England following the referendum.

2 Synthetic control method

The SCM was introduced by Abadie and Gardeazabal (2003), Abadie et al. (2010) and (2015) to address the difficulty in finding the counterfactual development of a treated unit. In general, the SCM assigns weights to control units so that these units best fit the pre-treatment characteristics of the treated unit. The SCM has been used in many fields, including international finance (Jinjarak et al. 2013; Sanso-Navarro 2011), financial policy (Aregger et al. 2017; Bruha and Tonner 2017; Opatrny 2017), trade liberalization (Billmeier and Nannicini 2013; Gathani et al. 2013; Hosny 2012) and taxation (Kleven et al. 2013).[2] Since the introduction of the SCM, there have been several articles that extend the SCM. For example, Acemoglu et al. (2016) and Cavallo et al. (2013) modify the SCM in such a way that more than one treated unit can be used to assess the intervention effect.

Another extension was proposed by Wong (2015), where the SCM is applied to a cross sectional setting and the synthetic control asymptotic distribution is derived as the number of individuals in the sample goes to infinity. Kreif et al. (2016) examine the SCM in contrast with the difference-in-differences method in the health policy context. They find that in contrast to the DiD method, for the incentivised condition, the SCM reports that a pay-for-performance (P4P) initiative did not significantly reduce mortality. Recently, Amjad et al. (2018), reviewed by Alberto Abadie, present a robust generalization of the synthetic control method for comparative case studies that automatically identifies a good subset of donors for the synthetic control, overcomes the challenges of missing data, and continues to work well in settings where covariate information may not be provided. Other development of the SCM was made by Ben-Michael et al. (2018). They propose Augmented SCM that extend SCM to settings where the fit on pretreatment outcome is infeasible. Precisely, they use the outcome model to estimate the bias due to poor pre-treatment

[2] See Firpo and Possebom (2017) for a rich list of studies using the SCM.

fit and then de-bias the original SCM estimate. We use this method as a robustness check for the original SCM.

The extent of inference procedures, originally developed by Abadie et al. (2010) and Abadie et al. (2015), represent an important research topic. Their inference procedures consist of estimating p-values through permutation tests. Using this procedure, they test the null hypothesis of no effect of the intervention. Ando and Ando (2015) design two new test statistics that have more power when applied to test the null hypothesis than those introduced by Abadie et al. (2010) and Abadie et al. (2015).

Another inference procedure that uses confidence intervals was proposed by Gobillon and Magnac (2016). They use a bootstrap technique to compute confidence intervals for the policy effect on more than one treated unit. To obtain valid results, a large number of treated and control regions is necessary. The issue with the validity of confidence intervals for a small number of control units was solved by Firpo and Possebom (2017). They extend the original inference procedures in a way that allows for different treatment assignment probabilities across units—any region could have a different probability of facing the intervention of interest. Moreover, their modified inference procedure allows for testing any type of sharp null hypothesis—any other than the null hypothesis of no effect proposed by Abadie et al. (2010) and Abadie et al. (2015). Finally, their inference procedure allows for the construction of confidence intervals for the post-intervention outcome as a function of time. We use the modified inference method of Firpo and Possebom (2017) to show the effect of the Brexit vote on long-term treasury yield (10-year gilt yield), Financial Times Stock Exchange 100 Index (FTSE 100 Index) and the real effective exchange rate (REER).

The following section is subdivided into three parts. The first presents the data used for the analysis, while the second and third describe the synthetic control method and its inference procedure, respectively. The notation and ideas mainly follow those of Abadie et al. (2010, 2015) and, for the extended inference procedure, those of Firpo and Possebom (2017).

2.1 Non-European OECD countries as control units

The data set used to analyse the impact of the Brexit vote on the selected variables is based on the OECD database. We use monthly data starting in April 2008 and ending in November 2018.[3] The donor pool of countries consists from 9 non-European OECD members—Australia, Canada, Chile, Israel, Japan, South Korea, Mexico, New Zealand, and the United States. As Abadie et al. (2015) suggest, countries that may be affected by intervention in the 'treated' country should be excluded from the sample. Therefore, we exclude from the donor pool European countries that could be affected the most by currency depreciation after the Brexit vote. Table 1 shows the overview of the total trade of OECD countries that is with the UK. Given the fact that the USA has the higher share of total trade with the UK, we provide additional

[3] The start period was chosen due to missing data.

results without the USA in the donor pool for REER and 10-Year gilt yield. We do not exclude the USA in the case of Stock Index due to low weight assigned by the SCM. Moreover, we provide the result for all OECD Countries for which we have full data (see Table 7 for reference).

We select long-term treasury yield (10-year gilt yield), Financial Times Stock Exchange 100 Index (FTSE 100 Index) and the real effective exchange rate (REER) as our outcome variables. Covariates include short-term treasury yield (2-year bond yield), 3-month interbank rate, overnight interbank rate (these variables help us to cover the short-term market perception of the current state), adjusted leading indicator, GDP ratio to trend, industrial production index, import and export indices, relative consumer price index, and harmonised unemployment (these variables explain the current state of economies in each country in the donor pool).[4] Finally, the selected covariates reasonably reflect the national monetary and financial sector, as well as the macroeconomic development of the economy.

2.2 Methodology

Suppose that the data set consists of information about $J + 1$ countries. Without loss of generality, we us assume that only the first country continuously faces the intervention of interest from period $t_0 \in \{1, \dots, T\}$ Abadie et al. (2010). As a consequence, there are J countries remaining as eventual control units that are not influenced by the intervention. Let Y_{it}^N denote the potential outcome of interest in the absence of the intervention for country i in period t, where $i \in \{1, \dots, J + 1\}$ and $t \in \{1, \dots, T\}$. Let T_0, where $1 \le T_0 \le T$, be the number of pre-intervention periods. Depending on the anticipation effect, we can reset T_0 to the period when the first effect of the intervention is assumed to appear Abadie et al. (2015). Let Y_{it}^I denote the outcome of interest affected by the intervention for country i in period $t \in \{1, \dots, T\}$. Naturally, we assume that the intervention has no effect on the outcome in the pre-intervention periods; therefore, $Y_{it}^N = Y_{it}^I$ for $t \in \{1, \dots, T_0\}$ (Opatrny 2017).

The effect of the intervention with $t > T_0$ is represented as follows:

$$\upsilon_{it} = Y_{it}^I - Y_{it}^N \tag{1}$$

Because Y_{it}^I is observed in Eq. (1), we must estimate Y_{it}^N. Abadie and Gardeazabal (2003) defined the weighted average of the control units with weights $w = \{w_2, \dots, w_{J+1}\}$ with $0 \le w_j \le 1$ for $j = 2, \dots, J + 1$ and

$$\sum_{j=2}^{J+1} w_j = 1$$

[4] See Table 2 for descriptive statistics of the variables.

These restrictions are made to avoid an extrapolation. By means of the given weights $\{w_2, \dots, w_{J+1}\}$, the synthetic control estimators of Y_{it}^N and υ_{it} are:[5]

$$\hat{Y}_{it}^N = w_2 Y_{2t} + \cdots + w_{J+1} Y_{J+1,t}$$
$$\hat{\upsilon}_{it} = Y_{it}^I - \hat{Y}_{it}^N$$

The next step is to choose the weights $\{w_2, \dots, w_{J+1}\}$ that best reflect the pre-intervention characteristics of the treated unit. Abadie et al. (2010) choose $w^* = \{w_2^*, \dots, w_{J+1}^*\}$, which minimizes:

$$v_1(X_{11} - w_2 X_{12} - \cdots - w_{J+1} X_{1,J+1})^2 + \cdots + v_k(X_{k1} - w_2 X_{k2} - \cdots - w_{J+1} X_{k,J+1})^2 \tag{2}$$

where $\{v_1, \dots, v_k\}$ represents the relative importance of the synthetic control assigned to predictors $\{X_{11}, \dots, X_{k,J+1}\}$. Consequently, the procedure comes down to choosing $\{v_1, \dots, v_k\}$.[6] As in most empirical studies using the SCM, we choose the weights $\{v_1, \dots, v_k\}$ to minimize the size of the prediction error, $Y_{it}^I - \hat{Y}_{it}^N$, in a selected pre-intervention period (Opatrny 2017).[7] This can be done by solving a nested optimization problem with v selected such that w minimizes the root mean square predicted error (RMSPE) during a selected period.[8]. Therefore, each choice of v results in a different country weight $w(v)$, which then gives a value for the RMSPE.

To precisely minimize the RMSPE, the following conditions must be fulfilled. First, countries that faced similar intervention should be excluded from the data set to avoid potential bias in the output. Second, to ensure a good fit of the counterfactual outcome, the control units must have similar economic performance to that of the unit exposed to the intervention. Moreover, countries that may be affected by the intervention in the 'treated' country should be excluded from the sample Abadie et al. (2015). Taking these assumptions into account, we consider non-European OECD countries as suitable comparison units. However, as Amjad et al. (2018) claim, for the optimal result, the method still depends on a subjectively chosen subset of donors and covariate matrix. Moreover, the SCM performs poorly in the case of missing data or strong levels of noise in the data set.

2.3 Inference procedures

We use three inferential methods for the SCM. Two of these methods are based on the 'placebo' effect initially introduced by Abadie and Gardeazabal (2003). The third method, which involves constructing a confidence interval, was briefly used in Opatrny (2017), but a later study Firpo and Possebom (2017) provides a theoretical background for constructing the confidence intervals for SCM.

5 See Abadie et al. (2010), where it is proved that $\hat{\upsilon}_{it}$ is an unbiased estimator of υ_{it}.

6 See Table 8 in Appendix for the relative importance of our covariates.

7 See Abadie et al. (2011), which describes other approaches for choosing the weights $\{v_1, \dots, v_k\}$

8 The RMSPE has the following formula: $RMSPE = \left(\frac{1}{T_0} \sum_{t=1}^{T_0} (Y_{1t} - \sum_{j=2}^{J+1} w_j^* Y_{jt})^2 \right)^{\frac{1}{2}}$

The first inference method applies the synthetic control method to all control units. As a result, we obtain a synthetic control for countries not exposed to the intervention. Consequently, this allows us to evaluate the estimation of the effect between the treated unit and the units not exposed to the intervention Opatrny (2017). In other words, we would lose confidence about the results if the SCM was used to estimate a large effect on a unit where the intervention did not occur. Formally, for each country $i \in \{1, \dots, J+1\}$ and period $t \in \{T_0, \dots, T\}$, Abadie et al. (2015) compare the effect of the intervention in the treated country, $\hat{\upsilon}_{1t}$, with the effect of the intervention in the control units $\hat{\upsilon}_{it}$. To solve the problem that $|\hat{\upsilon}_{1t}|$ could be atypically larger than $|\hat{\upsilon}_{it}|$ for some periods but not for others, they suggest using the distribution of the following statistic:

$$RMSPE_i := \frac{\sum_{t=T_0+1}^{T}(Y_{it} - \hat{Y_{it}^{N}})^2/(T-T_0)}{\sum_{t=1}^{T_0}(Y_{it} - \hat{Y_{it}^{N}})^2/(T_0)} \tag{3}$$

Equation 3 is then used to compute the following p-value:

$$p := \frac{\sum_{i=1}^{J+1} D_i}{J+1}, \tag{4}$$

where D_i equals 1 if ($RMSPE_i \geq RMSPE_1$). Therefore, Abadie et al. (2015) could reject the null hypothesis of no effect of the intervention if p is less than some prespecified significance level. However, Firpo and Possebom (2017) claim that the design of the p-value in Eq. 4 implicitly assumes a uniform distribution of the probability of being treated. Therefore, their extension of the inference method suggests a parametric form of treatment probabilities. For $\bar{i} \in \Omega := \{(1), \dots, (J+1)\}$, such that $RMSPE_{(1)} > RMSPE_{(2)} > \cdots > RMSPE_{(J+1)}$ and $RMSPE_{\bar{i}} = RMSPE^{obs}$, if there is more than one $i' \in \Omega$ with that property, Firpo and Possebom (2017) propose selecting the largest. They define the treatment probabilities as

$$\pi_{(i)}(\phi) = \frac{exp(\phi v_{(i)})}{\sum_{i' \in \Omega} exp(\phi v_{i'})}, \tag{5}$$

where $\phi \in R_+$ is the sensitivity parameter and $v_{i'} \in \{0, 1\}$ for each $i' \in \Omega$. This result provides an intuitive way to analyse the sensitivity of the parameter to deviations from the uniform distribution assumption. For example, the interpretation of ϕ is as follows: a unit $i_{(1)} \in \Omega$ with $v_{(i1)} = 1$ has a $\Phi := exp(\phi)$ times higher probability to be treated than a unit $i_{(2)} \in \Omega$ with $v_{(i2)} = 0$ (Firpo and Possebom 2017).[9] Due to assumption 5, the authors use the following formula to compute the p-value:

$$p(\phi, v) := \sum_{(i) \in \Omega} \frac{exp(\phi v_{(i)})}{\sum_{i' \in \Omega} exp(\phi v_{i'})} D_i, \tag{6}$$

[9] See section 3 in Firpo and Possebom (2017) for the details.

where D_i equals 1 if $(RMSPE_{(i)} \geq RMSPE_{\bar{i}})$ and $v := (v_1, \dots, v_{J+1})$. This formula allows us to reject the exact null hypothesis if $p(\phi, v)$ is less than some pre-specified significance level.

In the empirical section below, we use the approach of Firpo and Possebom (2017). Given the fact that $(J + 1) = 10$, the probability that any control unit would receive the same treatment effect reaches a maximum of 1/10, which is equivalent to a p-value of 0.1 according to Eq. 4 proposed by Abadie et al. (2015). Our results suggest p-values equal to 1/10, 3/10 and 6/10 for the 10-year gilt yield, real effective exchange rate and FTSE 100 Index, respectively. Applying the standard rejection rule when the p-value equals 0.1, we do not reject the exact null hypothesis H_0: There is no effect of the event $Y^N_{1t} = Y^I_{1t}$ for $t \in \{1, \dots, T\}$ in terms of REER and the FTSE 100 Index. Applying the sensitivity analysis proposed by Firpo and Possebom (2017) to check the robustness of the result for REER, we do not reject H_0: There is no effect of the event is not a robust result, because we must set $\phi_{REER} = 1.35$ in order to reject it at the 10% significance level. As Firpo and Possebom (2017) suggest, when the exact null hypothesis, H_0, is false and we do not reject it, we want the sensitivity parameter $\phi \in R_+$ to be small because a more robust result could keep us from making a type II error. We argue that $\phi_{REER} = 1.35$ is reasonably small according to section 5.2 in Firpo and Possebom (2017). On the other hand, in the case of the FTSE 100 Index, we must set $\phi_{FTSE_{100}} = 2.6$ in order to reject H_0 at the 10% significance level. When the exact null hypothesis, H_0, is true and we do not reject it, we want the sensitivity parameter $\phi \in R_+$ to be large because a more robust result could keep us from making a type I error. Finally, in the case of the 10-year gilt yield, we already reach a maximum p-value of 0.1, which results in $\phi_{10YBondYield} = 0.003$. As Firpo and Possebom (2017) claim, if the sensitivity parameter $\phi \in R_+$ is close to zero, the permutation test's decision is not robust to small violations of the assumption of a uniform distribution of the probability of being treated, i.e., $\phi = 0$.

In conclusion, our results indicate that H_0: There is no effect of the intervention, $Y^N_{1t} = Y^I_{1t}$ for $t \in \{1, \dots, T\}$, may be true for the FTSE 100 Index but false for the REER and 10-year gilt yield. This result is confirmed by the inference method using confidence sets described in the Synthetic Outcome Sect. 3 below.

The second method related to the placebo study applies the synthetic control method to the period when the intervention did not occur in a treated unit. As Abadie et al. (2015) mention, a large placebo estimate would undermine the credibility of the result. For example, if there is a significant effect of the intervention in an earlier period, the confidence of the effect would greatly diminish.[10]

The third method is based on the construction of a confidence interval. As mentioned earlier, in the study conducted by Opatrny (2017), point-wise confidence intervals are used. Using the original RMSPE[11] computed by the SCM, we derive the respective confidence sets for the outcome Y^N_{1t} in the post-intervention period $t \in \{T_0, \dots, T\}$. In this empirical research, we use the confidence sets proposed by Firpo and Possebom (2017), which provide the theoretical background for

[10] We can choose random periods prior to the intervention.

[11] The formula mentioned in the footnote in Sect. 2.2.

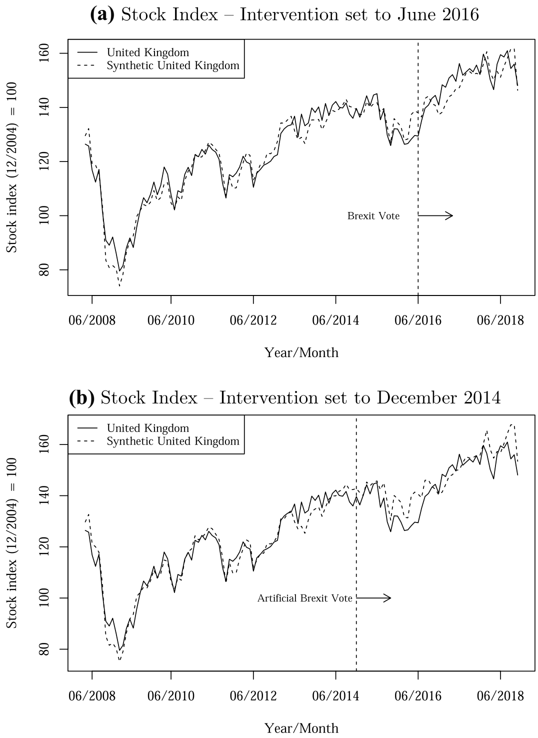


Fig. 1 Brexit vote had no impact on the Stock Index

confidence sets with constant and linear in time intervention effects. For linear in time intervention effect, they assume

$$H_0^{'} : Y_{it}^I = Y_{it}^N + (\hat{c} \times (t - T_0))D_t, \tag{7}$$

for each unit $i \in \{1, \ldots, J+1\}$ and time period $t \in \{1, \ldots, T\}$, where D_t equals 1 if $t \geq T_0 + 1$ and $\hat{c} \in R$.[12] Therefore, Firpo and Possebom (2017) assume constant in space effects, but linear in time intervention effects. Moreover, they suggest that we can apply the inference procedure described earlier in this Sect. 2.3 to the empirical distribution of $RMSPE^{\hat{c}}$ as a test statistic.[13]Consequently, the $(1 - \gamma)$—the confidence interval for the linear in time intervention effects—becomes

$$CI_{(1-\gamma)}(\phi, v) := \left\{ \begin{array}{l} f \in R^{\{1,\ldots,T\}} : f(t) = (RMPSE^{\hat{c}} \times (t - T_0)) * D_t \\ \text{and } p^{\hat{c}}(\phi) > \gamma \end{array} \right\} \subseteq CI_{(1-\gamma)}(\phi, v), \tag{8}$$

where $\gamma \in (0, 1) \subset R$. Intuitively, as Firpo and Possebom (2017) state, the confidence interval contains all linear in time intervention effects, for which $H_0^{'}$ is not rejected by the inference procedure described earlier in this Sect. 2.3.

[12] For constant in time intervention effects, they exclude the term $(t - T_0)$ from Eq. 7.

[13] The inference procedure is referred to as *the first method*.

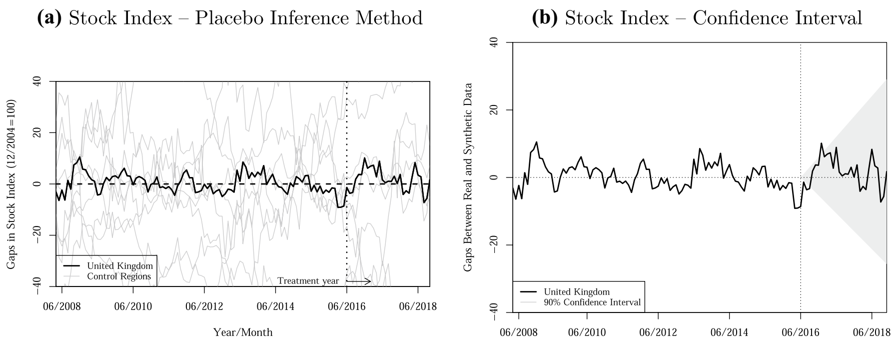


Fig. 2 Statistical Significance of the Synthetic Output for the Stock Index

3 Results

In Fig. 1a, we can see that the synthetic output almost copies the path of the real output after the Brexit vote. In other words, the FTSE 100 index would not have changed had there not been a referendum, taking into account the monthly frequency of the data.[14] The results are driven mainly by New Zealand, Mexico and the United States, with weights of 0.477, 0.125 and 0.103, respectively (Table 3 in Appendix). All other control units receive a weight less than 0.1. As Manning (2016) comments, equity markets declined remarkably immediately after the Brexit vote; however, they rebounded strongly in the following days. He adds that this recovery was driven in part by accommodative central bank policy and in part by 'search for yield' behaviour among market agents in an environment of low interest rates. In summary, our results suggest that the unexpected outcome of the Brexit vote had no impact on the equity market in the medium to long term.

As a robustness check of the results, we run the inference procedures described in Sect. 2.3. In Fig. 1b, we change the date of the vote to December 2014. The resulting dashed line is almost identical to that in Fig. 1a. Following the idea of Abadie et al. (2015), if there is not a significant effect of the Brexit vote in an earlier period, the confidence of the effect is not undermined. Our results do not undermine the main finding, as shown in Fig. 1b.

Figure 2 shows the statistical significance of the result. In Fig. 2a, we can see an estimate of the effect of the Brexit vote on the UK and all control units.[15] Moreover, the intervention effect does not abnormally differ from that of the other control units given the pre-treatment fit. Furthermore, we check whether changing the control group has any impact on the result. Therefore, we remove up to three controls from the donor pool with the highest weights. Figure 7 and Table 3 in Appendix show the results. In all cases, we get poor pre-intervention fit, which suggests that we cannot derive any significant inference from the robustness check. Alternatively,

[14] See Table 9 in Appendix for quantitative details about the result in the period after the Brexit vote.

[15] Additionally, we run SCM with all OECD countries in the donor pool. The results are not different from our original findings (Fig. 11b and Table 7).

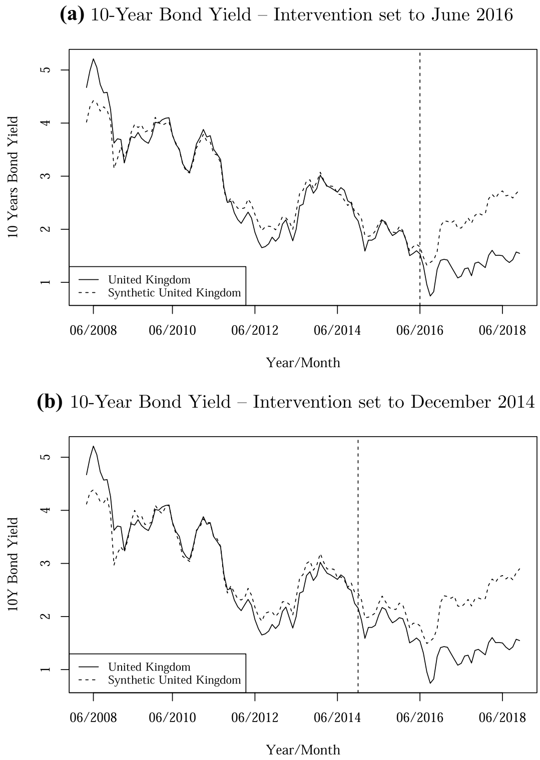


Fig. 3 Synthetic 10Y bond yield higher than the real one

we use the Augmented SCM to check the results (see Figs. 12b and 13b). These results confirm our original findings. Figure 2b shows the estimated 90% confidence intervals for the full control group. Intuitively, if the confidence interval includes the zero function, we do not reject the null hypothesis H_0': *There is no effect of the Brexit vote on the Stock Index*. However, immediately after the Brexit vote, the estimated result lies outside the confidence interval. This result is in line with the comments of Manning (2016) that equity markets significantly dropped after the Brexit vote but recovered quickly in the following period. In conclusion, the statistical inference procedure confirms the result that the Brexit vote had no impact on the Stock index in the medium to long term.

On the other hand, the impact of the Brexit vote on the 10-year gilt yield is depicted in Fig. 3. The synthetic output would have been 1.2 percentage points higher had there not been a referendum (Fig. 3a). The United States, Canada and South Korea compose the synthetic output with weights of 0.440, 0.332 and 0.226, respectively. Other countries receive weights of less than 0.002.[16]

In addition to the aforementioned hedging activities by pension funds and insurers, the low 10-year gilt yield after the Brexit vote was driven in part by increased expectation for an additional round of quantitative easing from the Bank of

[16] See Table 4 in Appendix for details.

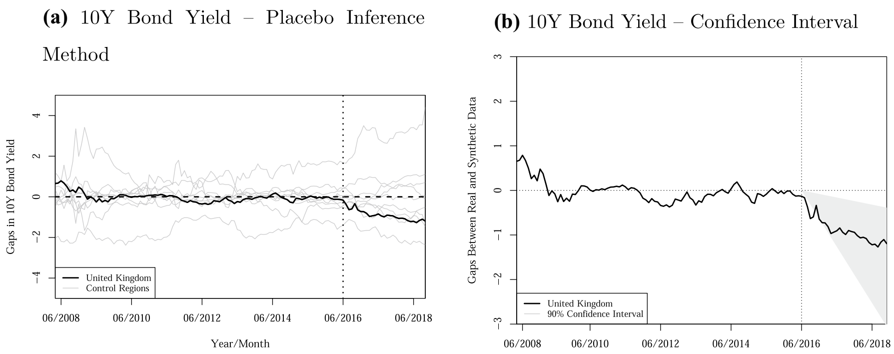


Fig. 4 Statistical significance of the synthetic output for long-term bonds

England. Consequently, on August 4, 2016, the UK Monetary Policy Committee (MPC) announced a supportive package for the UK economy consisting of (Manning (2016), p. 147): (1) a 25 basis points cut in Bank Rate to 0.25%, (2) a Term Funding Scheme (TFS) to reinforce the pass-through of the cut in Bank Rate, (3) purchases of up to £10 billion of sterling non-financial investment-grade corporate bonds (see 'Corporate capital markets' section), and (4) an increase in the stock of purchased UK government bonds, by £60 billion over six months, to £435 billion. These measures helped to further lower the 10-year gilt yield, as noted by Bank of England (2016) in Inflation Report. Another fact that explains the significant gap between the real and synthetic output is related to increased US Treasuries yield in the period after the Brexit vote. As Manning (2016) explains (pp. 148): 'market contacts attributes these moves to an increasing focus on domestic drivers of interest rates rather than spillovers from the United Kingdom's referendum and the subsequent monetary policy response'.

We verify the statistical significance by setting the Brexit vote to December 2014 (Fig. 3b). The result closely resembles that in Fig. 3a. Therefore, we believe that this statistical inference does not undermine our result. Alternatively, we run the placebo inference method (Fig. 4a). Given the result, we infer that the synthetic output does not abnormally differ from the other control units. In other words, applying the SCM on all other control units, we do not find any better pre-intervention fit.[17] To check the robustness of the control group we remove controls with the highest weights. Figure 8 and Table 4 show the result. We can see that the results in all cases relatively well resemble pre-intervention period in the UK. Moreover, the results in all cases suggest that there would have been higher yield had there not be a referendum. However, only the result without the USA is the most similar the one of the full control group.[18] The fact, that removing control units has relatively small impact on the pre-intervention fit, we provide another robustness check, where we shortened the

[17] Additionally, we run the SCM with all OECD countries in the donor pool. The result confirms our original findings (Fig. 11a and Table 7).

[18] Given this finding, Fig. 9a shows the result for statistical significance without the USA in the control group. The confidence interval includes the zero function, therefore the result is not statistically significant.

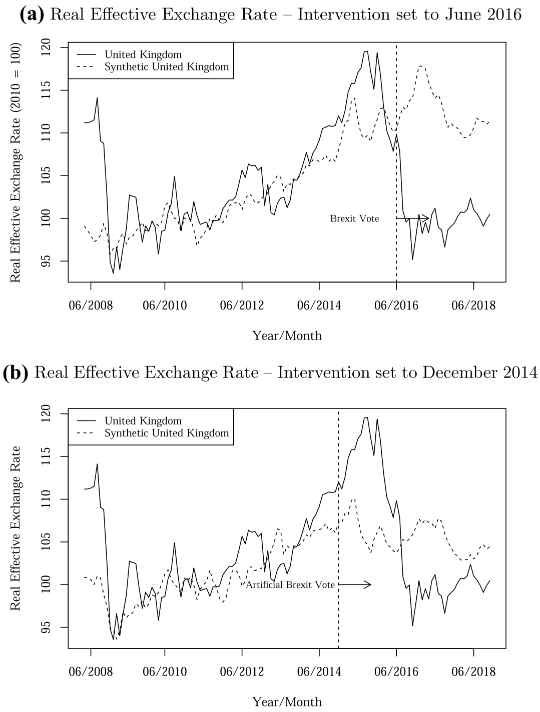


Fig. 5 Synthetic REER higher than the real one

pre-intervention period.[19] Figure 8b and Table 5 show the result for the period starting by January 2013. These results (see Fig. 9b) are almost identical to our original findings. Furthermore, we use the Augmented SCM which confirms our original findings (see Figs. 12a and 13a in Appendix) .

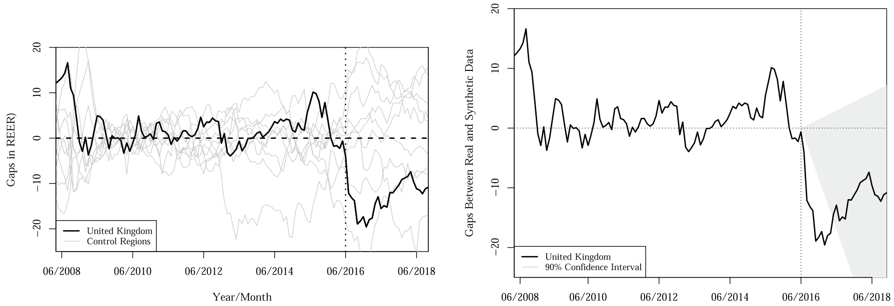
Finally, we estimate the 90% confidence interval (Fig. 4b). The confidence interval does not include the zero function; therefore, we reject the null hypothesis H_0': *There is no effect of the Brexit vote on 10-year gilt yield*. To conclude, we claim that the Brexit vote had a significant negative impact on the 10-year gilt yield.

Finally, we apply the SCM to the real effective exchange rate (Fig. 5a). The synthetic output would be 10.8 index points higher than the real one relative to the last period in our data set, November 2018.

The result is driven primarily by the United States, New Zealand and South Korea, with weights of 0.523, 0.388 and 0.054, respectively. Other control units receive weights of less than 0.004.[20] This is a rather intuitive result given the sharp depreciation in sterling right after the Brexit vote. As Manning (2016) mentions, sterling depreciated by approximately 7% on a trade-weighted basis. Moreover, he adds that the further depreciation in August could be attributed to the Bank of England's easing monetary policy. However, since then, the currency reversed the trend

[19] We thank to anonymous referee for this suggestion.

[20] See Table 6 in the Appendix for details.



(a) Real Effective Exchange Rate – Placebo Inference Method

(b) Real Effective Exchange Rate – Confidence Interval

Fig. 6 Statistical significance of the synthetic output

and gained back part of the loss. Consequently, Manning (2016) explains (p. 149): 'By the end of the review period (September 2016), forward-looking measures of implied volatility had decreased, and foreign exchange (FX) options pricing also pointed to a relatively even balance of upside and downside risks for the currency. However, speculative investors look to have remained more heavily positioned for the latter. Weekly positioning data for FX futures showed a continued increase in sterling short positions despite some improvement in UK economic data towards the end of the review period'. In conclusion, we see substantial short-term impact, which weakens in the long term.

In Fig. 5b, we set the Brexit vote to December 2014. The synthetic outcome follows the same trend as that in Fig. 5a. Therefore, this placebo inference does not undermine the result.

In Fig. 6a, we apply the SCM to all control units. The result does not abnormally differ from all other control units.[21] This suggests that we do not find any better pre-intervention fit among the control units. Alternatively, we provide results for various control units (see Fig. 10 and Table 6 in Appendix). These results reveals the similar effect as our original findings, however all of the results suffer from poor pre-intervention fit. Therefore, we cannot derive any significant inference. To overcome the problem with poor pre-intervention fit we employ the Augmented SCM (Figs. 12c and 13c). The results confirm our original findings. Finally, we estimate the 90% confidence interval (Fig. 6b).

The confidence interval does not include the REER result immediately after the Brexit vote. Therefore, there was a significant effect in the short term. Moreover, the confidence interval contains the zero function, which suggests that when considering a longer period, the REER would have been lower, but the result is not statistically significant.

[21] We run the SCM with all OECD countries in the donor pool. The results are not different from our original findings (Fig. 11c and Table 7).

4 Conclusion

We examine the impact of the Brexit vote on the UK stock exchange, long-term government bonds and the exchange rate. Using the synthetic control method developed by Abadie and Gardeazabal (2003), we build a counterfactual world that shows how selected variables would have developed had there not been a Brexit vote. We use Firpo and Possebom (2017)'s approach to assess the inference method from the SCM. Using well-developed quantitative analysis tools, this paper uniquely contributes to the current literature on the economic impact of the Brexit vote on the development of stock and bond markets and the real effective exchange rate.

Our results show that there would not have been any significant change in the development of the FTSE 100 Index in the mid to long term if the referendum had not occurred. On the other hand, we find a significantly negative effect of 1.2 percentage points on the 10-years bond yield. Given the geopolitical circumstances in mid 2016, financial agents investing in GBP could seek safer investment options represented by longer-term government bonds, which consequently could result in lower bond yields. Finally, we apply the SCM to the real effective exchange rate, which would have been 10.8 index points higher than the real one if the referendum had not occurred. However, these results could suffer from the main shortcomings of the method. As Amjad et al. (2018) note, the shortcomings are related to the fact that the optimal result depends on a subjectively chosen subset of donors and covariate matrix. Furthermore, the SCM performs poorly in the cases of missing data or strong levels of noise in the data set.

Overall, we estimated the effect of the Brexit vote on the UK's financial variables. Because the real terms and conditions of Brexit were still being negotiated during the writing of this paper, we could not estimate the economical impact of Brexit itself, which could be the subject of a future analysis.

Acknowledgements This project has received funding from GAUK No. 1250218 and from project SVV 260 463. The author is grateful to Tomas Havranek and Vaclav Broz for their valuable comments and suggestions.

Appendix 1: Descriptive statistics

See Tables 1 and 2.

Table 1 UK trade. *Source*: ONS UK Trade (data for 2016)

Country	% of total trade that is with the UK
Germany	10.70
Netherlands	6.42
France	6.37
Ireland	4.25
Spain	3.94
Belgium	3.61
Italy	3.55
Switzerland	3.07
Norway	1.81
Sweden	1.53
Denmark	0.94
Czech Republic	0.79
Portugal	0.69
Austria	0.62
Luxembourg	0.55
Greece	0.50
Hungary	0.45
Finland	0.44
Slovakia	0.30
Lithuania	0.14
Iceland	0.14
Latvia	0.10
Slovenia	0.07
Estonia	0.05
European OECD countries average	**2.10**
USA	15.20
Japan	2.03
Canada	1.46
Australia	1.28
Turkey	1.27
South Korea	1.02
Mexico	0.34
New Zealand	0.23
Chile	0.14
Non European OECD countries average	**2.55**

Table 2 Descriptive statistics of the variables used for the SCM computation. *Source*: Author's computation based on the OECD database

Statistic	N	Mean	SD	Min	Pctl(25)	Pctl(75)	Max
2YBonds	1234	2.292	1.977	0.050	0.569	3.490	8.870
10YBonds	1270	3.535	1.898	− 0.230	2.126	4.958	9.820
Index	1270	150.203	51.256	57.530	113.760	181.512	291.740
Value.3 month interbank rate	1233	2.185	1.953	0.050	0.550	3.190	8.870
Value.Leading indicator, amplitude adjusted	1270	99.814	1.333	93.508	99.387	100.564	103.474
Value.GDP (ratio to trend, smoothed)	1257	99.795	1.089	94.662	99.477	100.358	103.149
Value.Industrial production, s.a.	1246	98.219	6.438	63.959	94.994	101.967	118.358
Value.Overnight interbank rate	1268	1.879	1.698	− 0.071	0.428	2.990	8.260
Value.Relative consumer price indices	1270	102.428	12.849	75.946	93.930	108.336	154.476
Value.Labour Force Survey—quarterly rates Harmonised unemployment—monthly rates Total All persons	1224	5.553	1.626	2.200	4.248	6.625	10.231
Value.International Trade Imports Value (goods) Total	1270	3.861	16.895	− 44.429	− 5.526	14.171	64.682
Value.International Trade Exports Value (goods) Total	1270	3.564	16.561	− 43.678	− 6.528	13.166	54.423
REER	1270	99.090	10.212	67.847	94.862	105.561	123.368

Appendix 2: Robustness checks

Stock index robustness checks

See Fig. 7 and Table 3

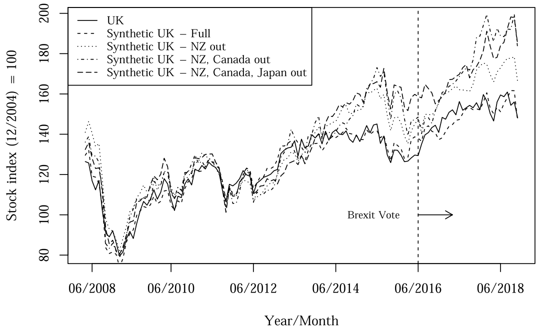


Fig. 7 Stock Index—changing control group. *Source*: Author's computation based on SCM

Table 3 Country weights computed by SCM—Stock Index. *Source*: Author's computation based on SCM

Country	Full sample	New Zealand out	New Zealand and Canada out	New Zealand, Canada and Japan out
Australia	0.027	0.257	0.060	0.000
Canada	0.047	0.481	–	–
Chile	0.050	0.001	0.159	0.004
Israel	0.032	0.007	0.051	0.369
Japan	0.080	0.242	0.643	–
Korea	0.060	0.003	0.005	0.000
Mexico	0.125	0.000	0.006	0.000
New Zealand	0.477	–	–	–
United States	0.103	0.008	0.078	0.627

10-year bond yield robustness checks

See Figs. 8, 9 and Tables 4, 5.

Fig. 8 Changing control group for 10-year bond yield results. *Source*: Author's computation based on SCM

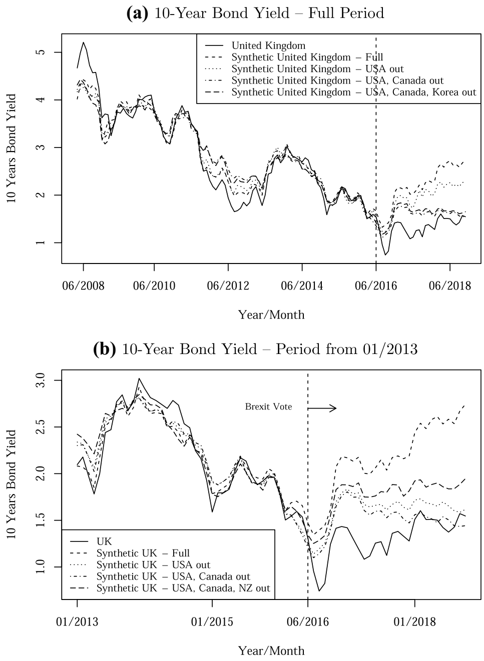


Fig. 9 Changing control group for 10-year bond yield results. *Source*: Author's computation based on SCM

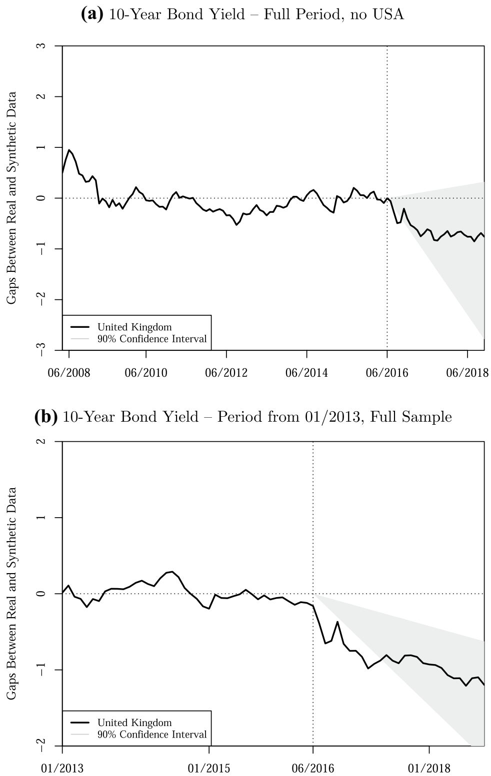


Table 4 Country Weights Computed by SCM—10Y Bonds. *Source*: Author's computation based on SCM

Country	Full sample	USA out	USA, Canada out	USA, Canada and Korea out
Australia	0.000	0.002	0.009	0.221
Canada	0.332	0.650	–	–
Chile	0.000	0.000	0.009	0.007
Israel	0.000	0.006	0.046	0.016
Japan	0.002	0.091	0.424	0.441
Korea	0.226	0.017	0.144	–
Mexico	0.000	0.000	0.005	0.017
New Zealand	0.000	0.233	0.364	0.297
United States	0.440	–	–	–

Table 5 Country weights computed by SCM—10Y Bonds, Period from 01/2013. *Source*: Author's computation based on SCM

Country	Full sample	USA out	USA, Canada out	USA, Canada and New Zealand out
Australia	0.042	0.000	0.000	0.512
Canada	0.251	0.008	–	–
Chile	0.008	0.001	0.001	0.000
Israel	0.011	0.005	0.000	0.000
Japan	0.075	0.454	0.479	0.427
Korea	0.008	0.109	0.003	0.003
Mexico	0.004	0.026	0.002	0.058
New Zealand	0.029	0.397	0.518	–
United States	0.572	–	–	–

REER robustness checks

See Fig. 10 and Table 6

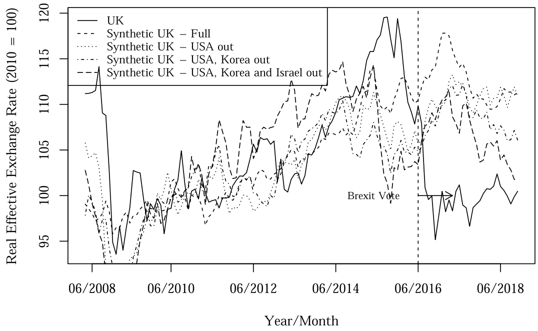


Fig. 10 REER—changing control group. *Source:* Author's computation based on SCM

Table 6 Country weights computed by SCM—REER. *Source*: Author's computation based on SCM

Country	Full sample	USA out	USA and Korea out	USA, Korea and Israel out
Australia	0.002	0.003	0.000	0.012
Canada	0.003	0.007	0.000	0.013
Chile	0.002	0.003	0.000	0.083
Israel	0.003	0.258	0.642	–
Japan	0.004	0.000	0.000	0.000
Korea	0.054	0.599	–	–
Mexico	0.021	0.003	0.000	0.056
New Zealand	0.388	0.125	0.358	0.836
United States	0.523	–	–	–

All OECD countries robustness check

See Fig. 11 and Table 7

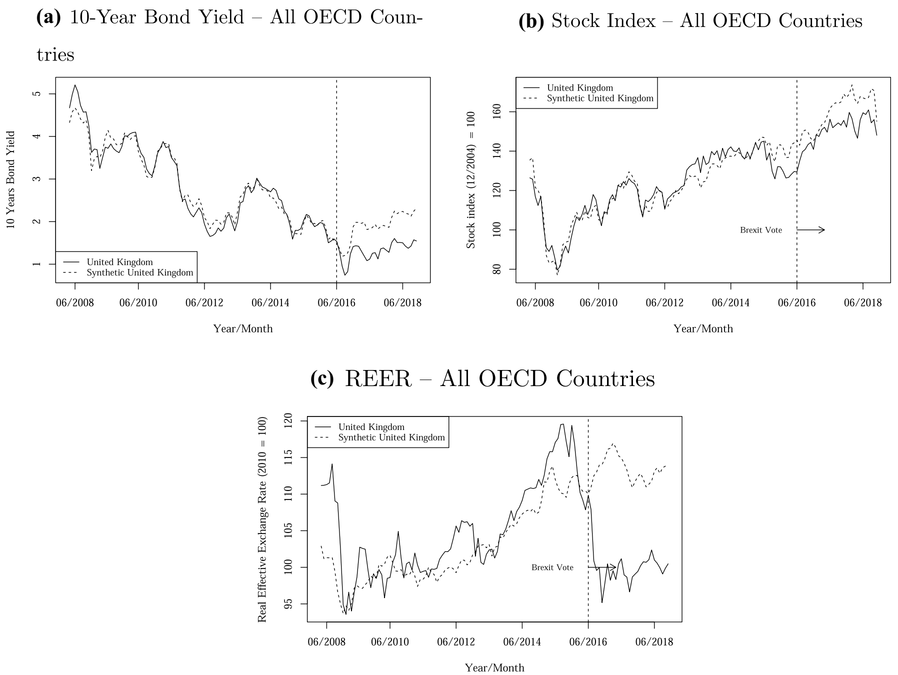


Fig. 11 10-year bond yield results. *Source*: Author's computation based on SCM

Table 7 Country weights computed by SCM—All OECD Countries. *Source*: Author's computation based on SCM

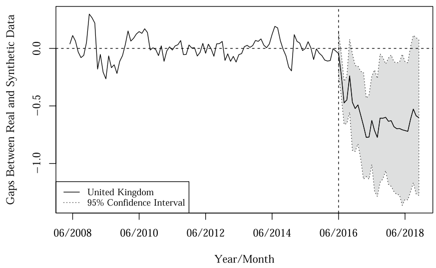
Country	10-year bond yield weights	Stock Index	REER
Australia	0.000	0.005	0.001
Austria	0.000	0.027	0.002
Belgium	0.000	0.016	0.001
Canada	0.000	0.005	0.000
Czech Republic	0.000	0.008	0.001
Denmark	0.000	0.015	0.001
Finland	0.311	0.048	0.003
France	0.000	0.012	0.001
Germany	0.000	0.014	0.001
Hungary	0.000	0.004	0.000
Chile	0.000	0.005	0.000
Ireland	0.000	0.009	0.001
Israel	0.000	0.012	0.001
Italy	0.000	0.014	0.001
Japan	0.000	0.000	0.000
Korea	0.000	0.234	0.389
Mexico	0.000	0.003	0.000
Netherlands	0.000	0.016	0.002
New Zealand	0.151	0.358	0.184
Norway	0.000	0.005	0.000
Poland	0.000	0.006	0.001
Portugal	0.000	0.009	0.001
Slovak Republic	0.000	0.021	0.001
Slovenia	0.000	0.010	0.001
Spain	0.000	0.015	0.001
Sweden	0.000	0.017	0.001
United States	0.536	0.085	0.405

Augmented SCM robustness check

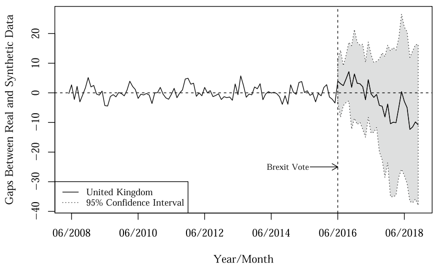
In all cases we use ridge regression as the outcome model with no covariates (Ben-Michael et al. 2018).

See Figs. 12 and 13

(a) 10-Year Bond Yield – All OECD Countries



(b) Stock Index – ASCM, All OECD Countries



(c) REER – ASCM, All OECD Countries

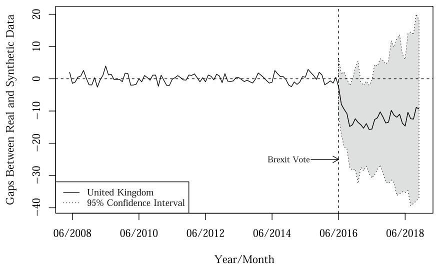
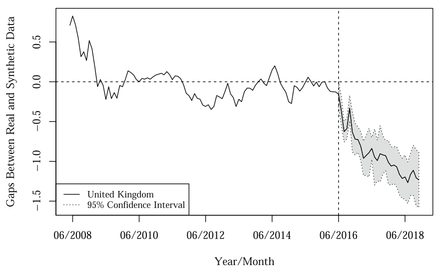
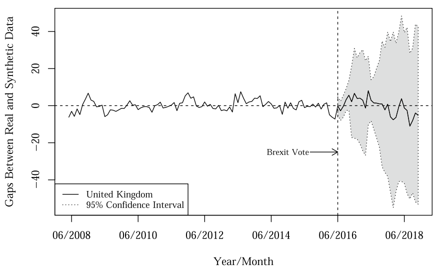


Fig. 12 ASCM, All OECD Countries. *Source*: Author's computation based on SCM

(a) 10-Year Bond Yield – Non - European OECD Countries



(b) Stock Index – ASCM, Non - European OECD Countries



(c) REER – ASCM, Non - European OECD Countries

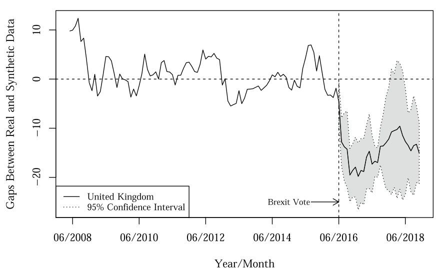


Fig. 13 ASCM, Non-European OECD countries. *Source*: Author's computation based on SCM

Appendix 3: Covariates weights

See Table 8.

Table 8 Weights for covariates computed by SCM. *Source*: Author's computation based on SCM

Statistic	Stock Index	10Y bonds	REER
2YBonds	0.0000	0.1176	0.0277
10YBonds	0.0001	0.0006	0.0365
Index	0.0311	0.0079	0.0450
Value.3 month interbank rate	0.0000	0.0798	0.0570
Value.Leading indicator, amplitude adjusted	0.1180	0.0676	0.0572
Value.GDP (ratio to trend, smoothed)	0.4549	0.0000	0.0450
Value.Industrial production, s.a.	0.0563	0.0000	0.0966
Value.Overnight interbank rate	0.0001	0.2003	0.1031
Value.Relative consumer price indices	0.0538	0.1707	0.1511
Value.Labour Force Survey—quarterly rates Harmonised unemployment—monthly rates Total All persons	0.0593	0.0602	0.1183
Value.International Trade Imports Value (goods) Total	0.0000	0.2508	0.0193
Value.International Trade Exports Value (goods) Total	0.0831	0.0245	0.0008
REER	0.1430	0.0198	0.2423

Numerical results

See Table 9

Table 9 Synthetic outcome results after the brexit vote. *Source*: Author's computation based on SCM

Period	Stock Index Real	Stock Index synthetic	10Y Bonds real	10Y Bonds synthetic	REER real	REER synthetic
6/2016	135.10	136.45	1.3105	1.4742	107.8845	111.8675
8/2016	139.68	143.32	0.9569	1.3250	100.8900	112.9908
9/2016	140.86	144.17	0.7421	1.3753	99.5689	112.6990
10/2016	143.31	141.44	0.8243	1.4192	99.9646	113.7787
11/2016	144.45	140.73	1.2430	1.5796	95.1671	114.0983
12/2016	140.91	137.29	1.4150	2.0581	97.7697	116.0484
01/2017	148.37	138.22	1.4336	2.1582	100.4659	117.7966
02/2017	147.46	141.32	1.4203	2.1494	98.2316	117.8141
03/2017	150.87	143.93	1.3058	2.1185	99.5233	117.5108
04/2017	152.11	144.79	1.1940	2.1621	98.3250	115.9445
05/2017	149.64	146.89	1.0830	2.0195	100.4665	114.8947
06/2017	156.20	147.28	1.1175	2.0241	101.1733	114.0848
07/2017	151.90	150.29	1.2560	2.0985	98.9577	114.4853
08/2017	153.13	152.66	1.2713	2.2113	98.6659	113.5466
09/2017	154.34	153.23	1.1241	2.1142	96.6353	111.8598
10/2017	153.14	152.18	1.3610	2.2611	98.6765	110.6707
11/2017	155.64	152.61	1.3816	2.3050	99.0702	111.1331
12/2017	152.19	152.06	1.3299	2.2707	99.4424	110.6367
01/2018	159.69	155.95	1.2781	2.2941	100.2237	110.5804
02/2018	156.48	160.58	1.5100	2.5756	100.7577	109.9158
03/2018	150.22	152.86	1.6034	2.6555	100.6829	109.4655
04/2018	146.58	150.17	1.5080	2.5874	101.0345	109.4961
05/2018	155.98	153.35	1.5108	2.6829	102.3698	109.7645
06/2018	159.49	151.15	1.5016	2.7221	101.0042	110.6012
07/2018	158.63	155.18	1.4229	2.6300	100.5650	111.7331
08/2018	160.95	158.15	1.3739	2.6439	99.9295	111.3530
09/2018	154.38	161.66	1.4260	2.5886	99.0835	111.3530
10/2018	156.00	161.60	1.5720	2.6747	99.8739	111.0281
11/2018	148.06	146.31	1.5462	2.7412	100.4784	111.3116

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