Nominal Shocks and Real Exchange Rates: Evidence from Two Centuries

William D. Craighead, Wesleyan University*

Pao-Lin Tien, Wesleyan University

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Abstract

This paper employs structural vector autoregression methods to examine the contribution of real and nominal shocks to real exchange rate movements using two hundred and seventeen years of data from Britain and the United States. Shocks are identified with long-run restrictions. The long time series makes possible an investigation of how the role of nominal shocks has evolved over time due to changes in the shock processes or to structural changes in the economy which might alter how a shock is transmitted to the real exchange rate. The sample is split at 1913, which is the end of the classical gold standard period, the last of the monetary regimes of the 19th century. The earlier subsample (1795-1913) shows a much stronger role for nominal shocks in explaining real exchange rate movements than the later subsample (1914-2010). Counterfactual analysis shows that the difference between the two periods is mainly due to the size of the nominal shocks rather than structural changes in the economy.

JEL Classification: F31, F41, N10

Key Words: *vector autoregression; monetary shocks; exchange rate movements; long-run identifying restrictions.*

*Corresponding author. Department of Economics, 238 Church St, Middletown CT 06459 USA. E-mail: bcraighead@wesleyan.edu. Tel.: 860-882-3845. Fax.: 860-685-2301.

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1. Introduction

Over the past two centuries, a number of institutional and structural changes have taken place that may have altered the behavior of real exchange rates. The 19th century economy suffered considerable financial instability, with recurrent banking panics and alternating periods of inflation and deflation. The 20th century saw a shift to fiat money and modern central banking, as well as changes in labor market institutions and in the degree of nominal rigidity present in the economy. While a number of previous studies have examined the extent to which real exchange rate fluctuations are driven by nominal or real shocks, this paper adds to the understanding of this question by investigating how the role of nominal shocks has changed over time.

In particular, this study examines the US-UK real exchange rate over the period 1794-2010 using vector autogressions (VARs) identified with long-run restrictions. Employing a data set that extends back to the late 18th century allows for an investigation of how changes in the shocks affecting the economy as well as in the structure of the economy may have influenced the role played by nominal shocks in real exchange rate behavior. Moreover, an improved understanding of the sources of movements in the US-UK real exchange rate may be of interest to scholars because this series has been widely utilized in exchange rate research, e.g. in studies of purchasing power parity by Lothian and Taylor (1996, 2008), among others.

While overall real exchange rate volatility increased in the 20th century, our estimation results indicate that nominal shocks have contributed more to real exchange rate movements in the period before 1913 than since. Results from a counterfactual experiment indicate that this may be attributable to a change in the relative size of nominal versus real shocks, rather than in the transmission of shocks to the real exchange rate. Although other studies have found evidence of changes in the relationship between prices and real quantities such as output and employment due to increased nominal rigidity over time, our results do not show a significant change in how nominal shocks are transmitted to the real exchange rate.

This research builds on a number of previous studies that have investigated the relative importance of real and nominal factors in real exchange rate dynamics using VARs identified with the long-run restriction methodology of Blanchard and Quah (1989). Using a two-variable system of real and nominal exchange rates between the US and five countries for 1973-89, Lastrapes (1992) finds that real shocks are far more important than nominal shocks. Clarida and Gali (1994) examine the real exchange rates between the US and Britain, Canada, Germany and Japan using a three-variable system motivated by an open-economy IS-LM model. They find that nominal shocks account for a substantial proportion of the variance of the Germany-US and Japan-US real exchange rates, but relatively little for Britain-US and Canada-US. Rogers (1999) considers a much longer time span, estimating a 5-variable system on US-UK data over the period 1889-1992. In his results, nominal shocks account for a much larger share of real exchange rate movements than Clarida and Gali found for the US-UK rate. Chen (2004) notes that because real exchange rates converge slowly, the identifying assumption that nominal shocks do not affect the real exchange rate in the long run can be problematic in shorter spans of data. He revisits Clarida and Gali's analysis using longer time samples – including the US-UK real exchange rate for 1889-1995 – and finds that doing so raises the proportion of real exchange rate variance attributed to nominal shocks. He also finds that results may be sensitive to choices of data series.

Other types of VAR identification approaches have also been used to examine real exchange rate behavior. Eichenbaum and Evans (1995) study the effect of monetary shocks on real exchange rates between the US and five other countries over 1974-1990 using a VAR with recursive short-run identifying restrictions. They find that monetary policy shocks generate persistent real exchange rate movements and deviations from uncovered interest parity. Although they find that monetary policy shocks account for a substantial portion of real exchange rate variation, the share attributable to monetary shocks is generally lower than what Rogers (1999) found. Kanas (2005) uses UK-US real exchange rate, interest rate and industrial production data for the period 1921-2002 to estimate a Markov-Switching VAR which allows for different parameters in high- and low-volatility regimes. He finds a much stronger link between real exchange rate movements and nominal shocks (interest rate differentials) than for real shocks (industrial production differentials). Muscatelli et al. (2007) estimate a four-variable system identified with short-run restrictions for the US-UK, US-Italy and Italy-UK real exchange rates for 1872-1998. They allow for the response to various shocks to change as the structure of the economy evolves by using time-varying parameters. For the US-UK real exchange rate, they find that productivity shocks have become more important in the postwar period and that the effect of monetary shocks was especially persistent during the Bretton Woods period. Fiscal shocks account for a substantial part of real exchange rate variation in the gold standard and interwar periods, but almost none during the post-Bretton Woods float.

Relevant historical literature related to potential sources of nominal shocks and changes in their transmission to real variables is discussed below in Section 2. The data employed in our empirical analysis is introduced in Section 3. Section 4 will detail the structural shock identification method, while Sections 5 and 6 outline the main estimation results and counterfactual analysis. Conclusions are offered in Section 7.

2. Historical Background

2.1 Sources of Nominal Instability

This subsection provides a brief overview of events such as banking crises and changes in monetary regimes that could have been sources of nominal shocks during the sample period. An accompanying chronology is provided in Table 1 for the first (1795-1913) and second (1914-2010) subsample periods we consider.

A fragile banking system contributed to monetary instability in the 19th century United States. Although the first (1791-1811) and second (1817-1836) Banks of the United States performed some central banking functions, it was not until December 1913 that the Federal Reserve Act was passed to establish a lender of last resort. In a chronology based on systematic examination of financial newspapers covering 1825-1929, Jalil (2012) identifies seven "major" banking panics[1], characterized by an increased demand for currency relative to bank deposits, and finds that these events had negative consequences for economic activity. Grossman (1993) finds that bank failures in the late 19th century US had significant macroeconomic impacts, primarily through a "nonmonetary" (disintermediation) channel. James et al. (in press) find that the bank suspensions of 1873, 1893 and 1907 were associated with macroeconomic downturns, which they attribute primarily to disruption of the payments system.

The US also saw a number of changes in the monetary standard in the 19th century. Initially, the US was on a bimetallic standard, however, whether silver or gold predominated depended on the ratio of silver to gold set in law – the "mint ratio" – relative to the market price ratio. The 15:1 ratio of the Coinage Act of 1792 undervalued gold at the mint, with the effect that the US was on a de facto silver standard. The ratio was revised in 1834 to 16:1, in effect shifting the US to a de facto gold standard[2]. Flandreau (1996) argues that the effect of bimetallism in France, which began in 1803, was to stabilize the relative price of silver to gold so it can be argued that US was effectively on the gold standard even earlier. The US issued paper money – "greenbacks" – during the Civil War and joined the gold standard at the beginning of 1879. However, the gold standard was controversial and deflationary conditions fueled a significant political movement for coinage of silver that culminated with William Jennings Bryan's 1896 presidential campaign. Uncertainty during this period about whether the gold standard would be maintained resulted in a risk premium on US debt, which was examined by Hallwood, MacDonald and Marsh (2000).

[1] In 1833-34, 1837, 1839, 1857, 1873, 1893 and 1907. Prior to the period covered by Jalil's study, US banks outside of New England suspended redemption of bank notes for specie (gold and silver) from August 1814 until February 1817 (Hepburn, 1903).

[2] The extent to which gold and silver circulated depended on a number of factors, including the degree to which the market price ratio deviated from the mint ratio, which varied over time; for a discussion see Martin (1968).

During the 20th century, the US suffered widespread bank failures in 1931-33 and left the gold standard in 1933. Under the Bretton Woods system that emerged following World War II, the dollar was again tied to gold. As US inflation rose in the 1960's, the system came under pressure and President Nixon announced the end of convertibility in August, 1971, and the US has been fully on a "fiat money" system since. The 1970's saw considerable inflation, but, following a the 1981-82 recession, which was brought on by a change in Fed policy after the 1979 appointment of Paul Volcker, US inflation has generally been modest.

The UK left the gold standard in 1797 when the Bank of England suspended gold payments after the threat of a French invasion sparked a banking panic. Convertibility of the pound into gold resumed in 1821. Although the UK did suffer several financial crises during the 19th century, it remained on the gold standard. While the Bank of England evolved a practice of acting as a lender of last resort during the 19th century[3], the UK also suffered periodic banking crises – notably in 1825, 1837, 1857, 1866 and 1890 – though these did not result in suspension of gold payments as they did in the United States.

During World War I, foreign exchange transactions were heavily restricted. After allowing the pound to float in 1919, the UK restored a fixed exchange rate at the prewar parity and joined the gold exchange standard in 1925. The deflationary implications of this are widely held to have contributed to the poor performance of the UK economy in the 1920's (e.g., Keynes 1925) and Britain suspended convertibility in September 1931. Again during World War II, foreign exchange trade was restricted. The controls were only gradually lifted following the war, with full convertibility restored in 1958 (see Yeager, 1976). The postwar Bretton Woods system allowed for discrete revaluations, which the UK availed itself of by devaluing the pound in 1949 and 1967. Inflation in UK during the 1970's was even more severe than in the US, and rapid depreciation of the pound led the UK to turn to the IMF for a loan in 1976. As in the US, inflation has generally been lower after a shift in monetary policy and a sharp recession in the early 1980's, though the UK did see a bout of inflation later in that decade. The UK joined the

[3] See Fetter (1965). Lovell (1957) and James (2012) provide evidence that the Bank of England was acting as a lender of last resort by the late 18th century, which is earlier than commonly believed.

European system of exchange rate bands known as the Exchange Rate Mechanism (ERM) in 1990, but was forced out by a crisis in September 1992. The Bank of England Act of 1998 granted the central bank independence and established an inflation targeting regime.

2.2 Changes in Nominal Rigidities

Although changes in and disruptions of the monetary system could be sources of nominal shocks, the impact on real variables would depend on the structure of the economy. In frameworks such as the Dornbusch (1976) "overshooting" model, nominal rigidities lead to higher real exchange rate volatility. Some evidence suggests that nominal rigidities increased over time, which could be a potentially important structural change affecting the propagation of nominal shocks to the real exchange rate.

One source of increasing nominal rigidities might have been changes in the labor market. In particular, wages may have become less flexible in the late 19th and early 20th centuries. For the US, Hanes (1993) found that the relationship between wage growth and the output gap decreased after the late 1880's and Hanes and James (2003) found that nominal wage cuts were common in 1841-1891, in contrast to the contemporary pattern where nominal wage changes are clustered at zero and cuts are quite rare (e.g., Akerlof, Dickens and Perry, 1996; Daly, Hobijn and Lucking, 2012). The institutional structure of British labor markets evolved substantially in the late 19th and early 20th century, with particularly rapid change in the period around the first World War. Thomas (1992) provides a discussion of this "institutional revolution" which included a substantial expansion of trade unions and collective bargaining. He finds a lack of sensitivity of British real wages to excess supply and, more broadly, that "the interwar labour market was, in a wide range of respects, characterized by inflexibility and inertia."

Studies at an aggregate level provide indirect evidence regarding nominal rigidities through the aggregate supply relationship between the price level and real output. Based on estimates from two-variable VARs for seven countries, Bayoumi and Eichengreen (1996) find that aggregate supply has changed from very steep in the classical gold standard period to relatively flat in the postwar period. They attribute this to increased inertia in prices. Using VARs to examine the gold standard period, Bordo, Landon-Lane and Redish (2009, 2010) find that, consistent with flexible prices, supply shocks account for most output variation while demand and monetary shocks had little impact on output in European economies (including the UK). However, they found that monetary shocks did have some effect on US output. For the US, James (1993) finds that the sensitivity of economic activity to monetary shocks increased earlier during the 19th century based on a comparison of VAR estimates for the 1836-57 and 1871-1909 periods. Beckworth (2007) employs a three-variable VAR on US data for 1864-97; he finds that "even though nominal rigidities may have been comparatively less binding during the postbellum period, they were still meaningful enough to influence real economic activity."

In a study directly focused on real exchange rates, Chernyshoff, Jacks and Taylor (2009) examine data for a panel of small countries in the classical gold standard and interwar periods. They find an increased relationship between terms of trade and real exchange rate volatility for countries on the gold standard in the interwar period. This is consistent with an increased role for nominal rigidities worldwide after World War I.

3. Data

The three-variable VAR model of Clarida and Gali (1994) is employed as the benchmark specification. The model includes the real exchange rate, real gross domestic product (GDP) and inflation rates for both the US and the UK. The variable definitions can be found in Table 2. For comparison purposes, the sample is split at 1913[4], which marks the end of the "classical" gold standard regime, the last of the 19th century monetary regimes. A significant institutional change

[4] The late 19th century was a period of transition and industrialization. We also consider an alternative break date of 1879, the beginning of the classical gold standard regime. This shifts more of this transitional period into the second subsample, but the results are qualitatively similar.

occurred in the US at the end of this year with the founding of the Federal Reserve, which Barsky et al. (1988) argue led to a worldwide change in the behavior of interest and prices.

Annual data on UK real GDP are from Broadberry et al. (2011) for 1794-1870, the Measuring Worth website for 1871-1947 and the UK Office of National Statistics for 1948-2010. US GDP is from Measuring Worth for 1794-1928 and the US Bureau of Economic Analysis for 1929-2010. The Measuring Worth GDP series are described in Officer and Williamson (2011a, b). Inflation is based on the deflators from the GDP series. The real exchange rate series are annual averages of data from Ahmad and Craighead (2011), which extends the series constructed by Craighead (2010) based on wholesale/producer price indexes.

In alternative specifications studied below as robustness checks, interest rate differentials as well as monetary base differentials are employed as monetary variables in place of the inflation differential. The US and UK nominal interest rates used are the consistent short-term rates from Measuring Worth (described in Officer, 2011). No short-term interest rate series is available for the US prior to 1830, so the Measuring Worth long-term rate series is adjusted by subtracting a term premium of 2.47 percent, which is the difference between long- and short-term rates in 1831. For 1794 through 1932, the monetary base series for both countries is from Officer (2002)[5]. Officer's dollar series is expressed in gold terms; it is converted, where appropriate, into "paper" dollar values. For 1933-1958, the US series is from Friedman and Schwartz (1963), and thereafter from the Federal Reserve. The UK monetary base series is from Capie and Weber (1985) for 1933-1982 and from the Bank of England thereafter. The UK and US both recently began paying interest on reserves, which considerably changed the relationship between reserves and money. For the period subsequent to these changes (2005 for the UK, 2007 for the US), data on currency is used as a proxy for the monetary base.

As a preliminary to estimating the VAR model, we establish some time series properties of the data. Table 3 reports results from unit root tests conducted for each variable in the

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[5] Officer's paper focuses on the US monetary base. We are grateful to him for supplying his UK monetary base series to us as well.

specifications considered. Because unit root tests are notorious for their lack of power, we perform both unit root and stationarity tests (ADF and KPSS) to ensure robustness.[6] In general, the test results conform to expectations. Relative output and the monetary base differential are non-stationary in levels and stationary in first-differences. The real exchange rate, inflation differential, and interest rate differential all appear to be stationary in levels for both subsamples. To facilitate comparison with Clarida and Gali's results and also to allow for permanent effects of demand shocks on it, the real exchange rate will enter the VAR in first differences.[7]

Cointegration tests are performed for the baseline specification as well as the various alternatives we consider to ensure that the models are not misspecified. The Johansen test results are somewhat sensitive to the lag order selection and test specification, but, in general, are supportive of the hypothesis of no cointegration among the three variables in the benchmark VAR as well as the robustness models. This finding is corroborated by the cointegration results reported in Chen (2004) who also uses the same Clarida and Gali specification, though with slightly different data and sample period.

4. Structural VAR Framework and Identification

The three-variable benchmark specification follows Clarida and Gali (1994). The variables of interest are contained in the vector $\mathbf{x} = \left[\Delta(y_{UK} - y_{US}), \Delta q, (\pi_{UK} - \pi_{US})\right]'$ which is assumed to follow a multivariate covariance stationary process and depend on lags of itself and some vector of structural shocks, $\boldsymbol{\varepsilon}$:

$$(1) \qquad A_0 \mathbf{x}_t = \sum_{j=1}^{k} A_j \mathbf{x}_{t-j} + \boldsymbol{\varepsilon}_t, \ \boldsymbol{\varepsilon}_t \sim N(0,D).$$

[6] ADF refers to the Augmented Dickey Fuller (1979) test with the null hypothesis of a unit root; and KPSS refers to the stationarity test developed by Kwiatkowski, Phillips, Schmidt, and Shin (1992), where the null hypothesis is stationarity (no unit root).

[7] A specification with the real exchange rate in levels was also examined. See Section 5.2 for details.

The structural shocks are assumed to be uncorrelated with each other so the variance-covariance matrix, D, is diagonal. Provided that the coefficient matrix A_0 is invertible, equation (1) can be rewritten more compactly as:

(2) $$A(L)\mathbf{x}_t = A_0^{-1}\boldsymbol{\varepsilon}_t,$$

where $A(L) = I - \sum_{j=1}^{k} A_0^{-1} A_j L^j$, and L is the lag operator. The Wold moving average representation of equation (2) is then

(3) $$\mathbf{x}_t = C(L)\boldsymbol{\varepsilon}_t,$$

where $C(L) = A(L)^{-1} A_0^{-1}$. The reduced-form Wold moving average representation of $\mathbf{x}$ is given by

(4) $$\mathbf{x}_t = E(L)\mathbf{v}_t, \quad \mathbf{v}_t \sim N(0,\Sigma),$$

where $E(L) = A(L)^{-1}$, and $\mathbf{v}_t = A_0^{-1}\boldsymbol{\varepsilon}_t$ is the vector of innovations such that each element in $\mathbf{v}$ is some linear combination of the structural shocks in $\boldsymbol{\varepsilon}$. The (reduced-form) autoregressive representation of the system in equation (4) can be given by

(5) $$A(L)\mathbf{x}_t = \mathbf{v}_t.$$

$A(L)$ can be consistently estimated using standard ordinary least squares (OLS). The residuals from the OLS regression can then be used to calculate Σ. The structural model, i.e., the coefficients of $C(L) = A(L)^{-1} A_0^{-1}$ will be identified to the extent that there are enough restrictions to determine the elements of $C(L)$ uniquely. In the case of long-run restrictions, as Blanchard and Quah (1989) established, by making assumptions about the long-run behavior of the variables in the model that will render *C(1)* to be lower triangular, one can invoke the relationship that the spectral density for **x** at frequency zero is proportional to the long-run variance-covariance matrix denoted Λ:

(6) $$\Lambda = E(1)\Sigma E(1)' = C(1)DC(1)',$$

such that a Cholesky decomposition of Λ provides a unique lower-triangular matrix that is equivalent to $C(1)D^{1/2}$. Given $C(1)$ and $A(1)$, the impact matrix, A_0^{-1}, can be obtained, and the vector of structural shocks, $\boldsymbol{\varepsilon}$, can be recovered.

We can express the long-run identifying restrictions on the benchmark model as:

(7) $$x = C(1)\boldsymbol{\varepsilon} \Rightarrow \begin{bmatrix} \Delta(y_{UK} - y_{US}) \\ \Delta q \\ (\boldsymbol{\pi}_{UK} - \boldsymbol{\pi}_{US}) \end{bmatrix} = \begin{bmatrix} C_{11}(1) & 0 & 0 \\ C_{21}(1) & C_{22}(1) & 0 \\ C_{31}(1) & C_{32}(1) & C_{33}(1) \end{bmatrix} \begin{bmatrix} \boldsymbol{\varepsilon}^s \\ \boldsymbol{\varepsilon}^d \\ \boldsymbol{\varepsilon}^m \end{bmatrix}$$

The lower-triangularity of $C(1)$ can be justified in a straightforward manner. Output is only affected by supply side disturbances (ε^s), such as a technology shock, in the long run while the real exchange rate can be permanently affected by both supply and demand side (ε^d) factors. Note that ε^d here refers to real aggregate demand disturbances not related to money, such as preference or fiscal shocks. The inflation differential will react to supply, demand, and nominal (ε^m) shocks, but since the inflation differential enters the VAR in levels these shocks do not have a long run permanent impact on it.

While long-run identification procedures are popular, there are some concerns about the reliability of structural inferences as outlined in Faust and Leeper (1997). We address these issues by considering alternative specifications for robustness and also by constructing confidence intervals for the impulse-responses and variance decompositions with the more reliable bias-corrected bootstrap method proposed by Kilian (1998). This has been shown by Kilian and Chang (2000) to have superior coverage accuracy when compared with more common methods of constructing confidence intervals for impulse-responses, such as those of Runkle (1987) and Lütkepohl (1990).

5. Estimation Results

5.1 Benchmark Model

The reduced-form US-UK benchmark VARs for both subsamples are estimated using 6 lags.[8] Figures 1 and 2 show the estimated dynamic response of the variables in the VAR to a one-standard deviation realization of a particular structural shock for each of the two subsamples. The estimates have been transformed to reflect the effect of shocks on the levels of the variables rather than their growth rates. The impulse-response functions shown in both figures generally conform to predictions of macroeconomic theory.[9] For example, consider the nominal shock (ε^m): as shown in the third column in Figure 1, it is associated with an increase in the inflation differential, $\pi_{uk} - \pi_{us}$. This indicates an expansionary UK monetary shock relative to the US, which, as expected, leads to an immediate depreciation of the real exchange rate, q, and an increase in relative output, $y_{uk} - y_{us}$. Similarly, for the second subsample (Figure 2), we also see an increase in the inflation differential along with an increase in the real exchange rate and relative output following a nominal shock. These results give us confidence that the nominal shock is identified correctly.

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[8] Subsample 1 data is from 1795-1913, but with 6 lags the actual estimation sample period is 1801-1913. Similarly for subsample 2, data is from 1914-2010, but actual estimation sample period is 1920-2010. The lag order is chosen using the sequential modified likelihood ratio (LR) test, which was used by Rogers (1999). This test typically selects a higher lag order than the standard lag selection criterions such as the Akaike Information Criterion (AIC), the Bayesian Information Criterion (BIC), and the Hannan-Quinn Criterion (HQC). The higher lag order is preferred because there is not much dynamic movement in the generated impulse responses with 1 lag. The cross sample comparison results we report are qualitatively similar regardless of the lag order.

[9] The confidence intervals reported are 68 percent (approximately 1 standard deviation) bootstrap bands.

Values of the structural nominal shock, ε^m, over time are shown in Figure 3. Although the US and UK have similar monetary histories, some of the large values of ε^m correspond to points where they have diverged. In particular, the inflationary monetary expansion associated with the Civil War is visible as a large negative shock in the early 1860's, while the deflation that followed as the US retired the "greenbacks" appears as a positive shock. The UK's departure from the gold standard in 1931 followed by that of the US in 1933 are visible as positive and negative shocks, respectively. A large positive shock also corresponds with the UK's 1976 inflation crisis.

The impact of the three structural shocks on the real exchange rate across the two subsamples can be compared using variance decompositions, which report the share of the variance of the forecasting error attributable to each of the structural shocks at a given time horizon. Table 4 presents the variance decomposition results for the two subsamples. For the earlier (1795-1913) subsample period (top panel) the most important shock, explaining about 76 percent of the variance in the real exchange rate on impact, is the demand shock, ε^d. The supply shock (ε^s) accounts for a mere 0.28 percent, but the nominal shock (ε^m) has a fairly strong effect at about 24 percent. As the forecast horizon expands, the demand shock becomes more dominant. The supply shock also gains some explanatory power while the effect of the nominal shock declines (as it should by construction). However, the nominal shock is quite persistent; even ten years after the shock hits the economy, it can still explain about 7 percent of the variability in the real exchange rate, more so than the supply shock. This is in contrast to Clarida and Gali's (1994) finding that nominal shocks explain no more than 2.2 percent of the variability in the bilateral real exchange rate between the US and the UK. The results for the second subsample (1914-2010; second panel of Table 4) are more in line with the results reported by Clarida and Gali (1994). The demand shock is still dominant, even more so than in the first subsample. The supply shock effect is a little stronger as well. This comes at the cost of a reduction in the explanatory power of the nominal shock, which at the peak only explains about 5 percent of the variability in the real exchange rate (at the 1 year forecast horizon). However, the nominal shock is still quite persistent, explaining about 3 percent of the variability in the real exchange rate ten years after the shock impacts the economy.

Overall, the results in Table 4 show a clear contrast between the two sample periods. Prior to the end of the classical gold standard, nominal shocks explain a substantial portion of the variability in the real exchange rate while in the post-gold standard period nominal shocks contribute relatively little to real exchange rate movements. Whether this change in the importance of nominal shocks is due to a change in the size of the shocks or the structure of the economy will be investigated in Section 6.

5.2 Alternative Specifications

To ensure the robustness of our results, we briefly consider three alternative specifications. First, since the unit root test results appear to show the real exchange rate, q, to be stationary in levels for the two subsamples, we try a specification (Robustness Model 1) with q entering the VAR in levels. With long-run identifying restrictions, if a variable enters the VAR in levels, it implies that no shock can have a permanent impact on the value of that variable in the long-run. Also, the variable in levels needs to be ordered last for the structural shocks to have a natural economic interpretation. Hence the ordering of the three variables in this specification is:

$$(8) \qquad \mathbf{x} = C(1)\boldsymbol{\varepsilon} \Rightarrow \begin{bmatrix} \Delta(y_{UK} - y_{US}) \\ (\pi_{UK} - \pi_{US}) \\ q \end{bmatrix} = \begin{bmatrix} C_{11}(1) & 0 & 0 \\ C_{21}(1) & C_{22}(1) & 0 \\ C_{31}(1) & C_{32}(1) & C_{33}(1) \end{bmatrix} \begin{bmatrix} \varepsilon^{s} \\ \varepsilon^{m} \\ \varepsilon^{d} \end{bmatrix}$$

The supply shock (ε^s) has a long-run impact on output and inflation differentials. The nominal shock (ε^m) has long-run impact on inflation differential. And with q ordered last and in levels, all structural shocks can have short-term effects on q but none would leave a long-run impact (since q is stationary). The variance decomposition for the real exchange rate in this model is shown in

Table 5.[10] The findings from the baseline model above remain robust with the real exchange rate in levels. The first subsample shows a much stronger effect of nominal shocks on the real exchange rate – for example, on impact, about 38 percent of the variability in the real exchange rate can be accounted for by the nominal shock versus only about 11 percent for the second subsample.

The next two robustness checks involve employing alternative measures of monetary conditions. The inflation differential variable is replaced with the nominal interest rate differential (Robustness Model 2) and monetary base growth differential (Robustness Model 3). Tables 6 and 7 report the variance decomposition results for these two versions of the model.[11] The same patterns emerge for these two specifications as in the benchmark model. Nominal shocks appear to be much more important in explaining real exchange rate volatility in the first subsample period.

6. Counterfactual Analysis

In this section, we investigate whether the decrease in proportion of real exchange rate variability that can be attributed to nominal shocks is driven by a change in the shocks (smaller nominal shocks in the later period) or in the propagation mechanism (structural changes in the economy that reduced the sensitivity of real exchange rate to nominal shocks). To address this, we perform counterfactual experiments where we exchange the estimated covariance matrices between the two samples and see if that has a noticeable impact on real exchange rate variability. The literature on the "Great Moderation" of output volatility has used counterfactual experiments

[10] The model is estimated with 1 lag for both subsamples. Lag order is selected based on the BIC since the LR test does not give a conclusive result in this case. We only report variance decomposition results on impact of the shock and two years after because only short-term results can be considered in this model. For this reason, impulse-responses are omitted as well.

[11] Robustness Model 2 is estimated with 6 lags for both subsamples and Robustness Model 3 is estimated with 9 lags. Lag order is selected based on the LR test.

extensively to investigate the potential causes of the moderation (e.g. Stock and Watson, 2002; Ahmed, Levin, and Wilson, 2004; Kim, Morley, and Piger, 2008).

Kim, Morley, and Piger (2008) used the following intuitive example to illustrate the principles behind counterfactual experiments in the context of potential structural changes. Consider a variable y_t that can be expressed as a stationary AR(1) process:

$$(9) \qquad y_t = \phi y_{t-1} + u_t \qquad u_t \sim N(0, \sigma^2)$$

Since y_t is stationary, $|\phi| < 1$, and the variance of y_t can be constructed as:

$$(10) \qquad \operatorname{var}(y_t) \equiv \gamma = \frac{\sigma^2}{1-\phi^2}.$$

The variance of the shock is σ^2 and ϕ represents the propagation mechanism (the structure of y_t as specified by the AR(1) model). Now suppose y_t is subjected to a structural break that corresponds to an increase in its variance. The variance of y_t could be altered by changes in either σ^2 or ϕ or both. Counterfactual experiments allow us to consider a hypothetical situation where only the shock variance or only the propagation mechanism had changed and analyze the resulting hypothetical variance (or other properties of y_t that can be constructed using σ and ϕ, such as impulse-responses and variance decompositions in the multivariate case). To illustrate, let i index the structural regime for shocks and j the structural regime for propagation. If we split the sample into two subsamples, then we have two possible regimes, $i, j = 1, 2$. The variance of y_t is:

$$(11) \qquad \gamma^{i,j} = \frac{(\sigma^i)^2}{1-(\phi^j)^2}.$$

Note that $\gamma^{1,1}$ gives us the actual variance for the structural regime 1 (subsample 1); $\gamma^{2,2}$ gives us the actual variance for the structural regime 2 (subsample 2); while $\gamma^{1,2}$ or $\gamma^{2,1}$ give us the counterfactual variances based on changes in only the shock component or only the propagation mechanism.[12]

Conducting the type of counterfactual experiment described above with our benchmark model allows us to examine whether the increase in real exchange rate volatility over time is driven by changes in the shocks or in the propagation mechanism. The first two columns of Table 8 report the standard deviation of the change in the real exchange rate (first difference of the log real exchange rate, or Δq) for each of the sample periods considered. Sample 1 (1795-1913) has a standard deviation of 0.0676 compared with 0.0753 for sample 2 (1914-2010). Table 8 also reports the standard deviation of Δq for the two sample periods using the VAR estimates (columns 3 and 4).[13] The numbers match up fairly well with the sample data in the first two columns, though the reported standard deviation for sample 2 is slightly higher. The last two columns give the counterfactual results. The counterfactual standard deviation under the sample 1 propagation mechanism with sample 2 shocks (0.0806) has a similar magnitude to that estimated for sample 2 (0.0831). Applying sample 1 shocks to the sample 2 propagation mechanism yields a counterfactual standard deviation (0.0685) that has a similar magnitude to that reported for sample 1 (0.0665). This means that (i) if the shocks of the second period had occurred during the first sample period, the change in the real exchange rate from 1795-1913 would have had a much higher variance; (ii) if the shocks of the first period had occurred during the second sample period, the standard deviation of Δq would be lower from 1914-2010. This implies that a change in the shocks is the main reason for the increase in the observed volatility of real exchange rate. Note that the variance decomposition results reported earlier in Tables 4 indicate that the importance of monetary shocks has declined for the real exchange rate. This, in

[12] As detailed in Kim, Morley and Piger (2008), this type of experiment does not provide a formal decomposition of what caused the variance of y_t to change. Counterfactual variances represent hypothetical scenarios only, though they may shed some light on potential causes of the structural change. Hence some caution is warranted in interpreting counterfactual results.

[13] An increase in volatility also occurred for the level of the real exchange rate, q.

combination with the counterfactual result reported in Table 8, would signify that the increase in real exchange rate volatility is driven by more volatile real shocks, not nominal shocks.

The question of whether the reduced importance of monetary shocks for real exchange rate variability in subsample 2 is due to a decline in the size of monetary shocks during that period or a change in the shock propagation mechanism (due, for example, to an increase in nominal rigidity) can be considered using counterfactual impulse responses (Figure 4) and variance decompositions (Table 9). These are produced similarly to the above by exchanging the estimated variance-covariance matrices between the two samples while maintaining the values of the estimated VAR coefficients for each sample. As one would expect, the effect of a nominal shock on real exchange rate declines for 1795-1913 when subsample 2 shocks applied, and increases for 1914-2010 when subsample 1 shocks are used. The results here suggest that the decline in the proportion of real exchange rate variability that can be accounted for by nominal shocks can be explained by a decline in the size of the shocks rather than a change in propagation. In fact, if we calculate the variance of the structural nominal shock ε^m backed out of the estimated residuals, the variance for sample 1 is about 0.2845 while the variance for sample 2 is 0.2446.

The reduction in the size of nominal shocks is consistent with fewer severe disturbances from banking crises in the 20th century compared to the 19th combined with the view that these events had real effects. While both countries struggled with inflation in the 1970s, after the interwar period, the alternating periods of inflation and deflation that characterized the 19th century were absent. The reduction in monetary instability in the United States may partly be attributed to the establishment of the Federal Reserve in 1913 and, more broadly, to institutional changes that reduced the frequency of banking crises in both countries in the 20th century. Also, the second subsample encompasses two periods noted for their relative economic stability: the Bretton Woods era and the "great moderation" period.

7. Conclusion

Overall, it is a robust finding that nominal shocks are more important in explaining real exchange rate fluctuations in the sample period prior to the end of the classical gold standard (1795-1913) compared with the later period (1914-2010). The shares of real exchange rate variance attributable to nominal shocks reported in Tables 4-7 are well within the estimated range found in the literature using either long-run or non-long-run identification schemes. However, considering a longer sample of data, and breaking it into two potentially structurally different time periods, has allowed us to examine how the role of nominal shocks has changed. While our findings of a small role for nominal shocks in the 20th century is consistent with the results of Clarida and Gali (1994), we show that nominal shocks accounted for a much greater share of real exchange rate movements in the 19th century. Counterfactual experiments indicate that the decrease in the importance of nominal shocks for real exchange rate volatility after the end of the classical gold standard is primarily attributable to a decrease in the volatility of the shocks themselves rather than a change in the structure of the economy.

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TABLE 1

SELECTED US AND UK MONETARY EVENTS 1795-2010

Year	Description
Subsample 1:	
1797	Suspension of specie payments by Bank of England
1814-1816	UK bank failures
1814-1817	US bank suspsensions
1821	Resumption of specie payments by Bank of England
1825	UK banking crisis
1834	US Coinage Act of 1834 reduces gold content of US dollar
1837	UK banking crisis
1837-1842	Suspension of specie payments by US banks
1857	UK and US banking crisis
1862	Beginning of US “Greenback” paper money standard period
1866	UK banking crisis
1873	US banking crisis
1879	Resumption of gold standard in US
1890	UK banking crisis
1893	US banking crisis
1896	US election of 1896; “Cross of Gold” speech
1907	US banking crisis
1913	Passage of Federal Reserve Act in December
Subsample 2:	
1914	Beginning of wartime restrictions on payments
1919	End of wartime restrictions; UK pound sterling floats
1925	UK returns to gold standard at prewar parity
1931	UK leaves gold standard
1931-1933	US banking crisis
1933	US leaves gold standard
1939	Beginning of wartime restrictions on payments
1949	UK pound sterling devalued
1958	End of UK exchange controls
1967	UK pound sterling devalued
1971	US closes “gold window”
1976	UK inflation crisis; IMF program
1979	Shift in monetary policies in both US and UK
1990	UK joins ERM
1992	UK leaves ERM
1998	Inflation targeting established in UK
2008	Global financial crisis; run on “shadow banking” system

TABLE 2

VARIABLE DEFINITIONS AND DATA SOURCES

Variable	Definition and Sources
y_{UK} - y_{US}	Relative output = ln(UK real GDP) – ln(US real GDP) UK GDP: 1794-1870: Broadberry et al. (2011) 1871-1947: Officer and Williamson (2011a)/Measuring Worth 1948-2010: UK Office of National Statistics US GDP: 1794-1928: Officer and Williamson (2011b)/Measuring Worth 1929-2010: US Bureau of Economic Analysis
q	Real exchange rate = ln(nominal exchange rate) + ln(US wholesale prices) – ln(UK wholesale prices) Ahmad and Craighead (2011); annual averages; nominal exchange rate is UK Pounds per US Dollar
π_{UK} - π_{US}	Inflation differential = (% change in UK GDP deflator) – (% change in US GDP deflator) Deflators calculated from GDP data described above
i_{UK} - i_{US}	Nominal interest rate differential = UK nominal interest rate – US nominal interest rate Officer (2011)/Measuring Worth "consistent" short term rate for ordinary funds. For the US, no short term series is available before 1830, so the long term rate is used, adjusted by subtracting 2.47, which is the difference between the two series 1831 (i.e., the term premium)
MB_{UK} -MB_{US}	Nominal monetary base differential = ln(UK nominal monetary base) – ln(US nominal monetary base) UK monetary base: 1794-1932: Officer (2002) 1933-1982: Capie and Webber (1982) 1983-2010: Bank of England (M0 through 2005, Notes and Currency thereafter) US monetary base: 1794-1932: Officer (2002), adjusted for paper currency depreciation 1933-1958: Friedman and Schwartz (1963) 1959-2010: Federal Reserve (Monetary Base through 2001, Currency Component of M1 thereafter) All values are end-of-year/December

TABLE 3

UNIT ROOT AND STATIONARITY TEST RESULTS

Variable	Structure	# of lags	ADF statistic	KPSS statistic
Subsample 1: 1795 – 1913 (for benchmark model)				
y_{UK} - y_{US}	C + T	0	-3.4301[1] (non-stationary)	0.0723[4] (stationary)
q	C	8	-3.3928[2] (stationary)	0.2190[4] (stationary)
π_{UK} - π_{US}	C	3	-6.9098[3] (stationary)	0.0925[4] (stationary)
$\Delta(y_{UK}$ - $y_{US})$	C	0	-12.7115[3] (stationary)	0.0565[4] (stationary)
Δq	C	0	-9.7304[3] (stationary)	0.1076[4] (stationary)
MB_{UK} - MB_{US}	C + T	4	-3.1210[1] (non-stationary)	0.1506[5] (non-stationary)
$\Delta(MB_{UK}$ - $MB_{US})$	C	3	-6.5823[3] (stationary)	0.2478[4] (stationary)
Subample 1: 1798 – 1913 (for Robustness Model 2)				
y_{UK} - y_{US}	C + T	0	-3.3912[1] (non-stationary)	0.0791[4] (stationary)
q	C	1	-4.6930[3] (stationary)	0.2673[4] (stationary)
i_{UK} - i_{US}	C	1	-4.0291[3] (stationary)	0.5095[5] (non-stationary)
$\Delta(y_{UK}$ - $y_{US})$	C	0	-12.4996[3] (stationary)	0.0702[4] (stationary)
Δq	C	4	-6.3127[3] (stationary)	0.1227[4] (stationary)
Subsample 2: 1914 – 2010				
y_{UK} - y_{US}	C + T	7	-2.5571[1] (non-stationary)	0.2272[6] (non-stationary)
q	C	0	-3.0431[2] (stationary)	0.2414[4] (stationary)
π_{UK} - π_{US}	C	10	-1.8858[1] (non-stationary)	0.1138[4] (stationary)
i_{UK} - i_{US}	C	8	-2.4051[1] (non-stationary)	0.2496[4] (stationary)
$\Delta(y_{UK}$ - $y_{US})$	C	10	-4.8573[3] (stationary)	0.1646[4] (stationary)
Δq	C	4	-6.1255[3] (stationary)	0.0418[4] (stationary)
MB_{UK} - MB_{US}	C	8	-2.7652[1] (non-stationary)	0.2340[4] (stationary)
$\Delta(MB_{UK}$ - $MB_{US})$	C	3	-3.2229[2] (stationary)	0.1298[4] (stationary)

1. Null hypothesis of unit root cannot be rejected at 5% level.
2. Null hypothesis of unit root can be rejected at 5% level.
3. Null hypothesis of unit root can be rejected at 1% level.
4. Null hypothesis of stationarity cannot be rejected at 5% level.
5. Null hypothesis of stationarity can be rejected at 5% level.
6. Null hypothesis of stationarity can be rejected at 1% level.

Note: All calculations performed in Eviews. Structure (C refers to constant and T refers to trend) and lag number selection follow the procedure laid out in Campbell and Perron (1991) with a starting maximum lag order of 10. ADF refers to the Augmented Dickey Fuller (1979) test with the null hypothesis of unit root; and KPSS refers to the stationarity test developed by Kwiatkowski, Phillips, Schmidt, and Shin, where the null hypothesis is stationarity (or no unit root). Technical details of these tests can be found in the Eviews Help file under "Unit Root Tests." Lag selection for the KPSS test is done automatically in Eviews using the Newey-West Bandwith instead of user specified as in the ADF test.

TABLE 4

VARIANCE DECOMPOSITION RESULTS FOR *q*
(BENCHMARK MODEL)

Time Horizon	ε^s (%)	ε^d (%)	ε^m (%)
Subsample 1: 1795 – 1913			
0 Year	0.28 (0.17, 9.48)	75.50 (48.87, 88.57)	24.22 (6.29, 46.33)
1 Year	0.16 (0.46, 9.64)	81.23 (55.92, 90.86)	18.62 (4.79, 38.92)
2 Years	0.20 (0.76, 10.35)	85.04 (62.52, 91.55)	14.76 (3.97, 32.04)
5 Years	5.25 (2.97, 18.53)	85.39 (66.54, 88.75)	9.36 (3.83, 20.21)
10 Years	5.25 (3.49, 20.43)	87.44 (68.64, 89.51)	7.30 (3.09, 15.79)
Subsample 2: 1914 – 2010			
0 Year	2.74 (0.30, 11.66)	93.48 (70.25, 96.84)	3.79 (0.43, 21.62)
1 Year	4.56 (0.84, 15.00)	90.75 (66.93, 95.32)	4.69 (0.82, 21.52)
2 Years	4.06 (1.23, 15.00)	92.04 (68.93, 94.76)	3.91 (1.09, 18.99)
5 Years	3.59 (2.81, 16.74)	93.25 (71.03, 91.87)	3.15 (1.79, 15.58)
10 Years	8.10 (6.16, 24.03)	88.86 (66.36, 87.94)	3.03 (2.09, 13.48)

Note: Benchmark model is estimated with 6 lags. Confidence intervals are obtained using the Kilian (1998) bias-corrected bootstrap method. The lower band indicates the 16th percentile of the bootstrapped distribution. The upper band indicates the 84th percentile.

TABLE 5

VARIANCE DECOMPOSITION RESULTS FOR q
(ROBUSTNESS MODEL 1 WITH q IN LEVELS)

Time Horizon	ε^s (%)	ε^d (%)	ε^m (%)
Subsample 1: 1795 – 1913			
0 Year	9.29 (1.54, 30.82)	52.29 (29.77, 70.86)	38.43 (15.09, 52.95)
1 Year	10.35 (2.04, 32.84)	60.50 (37.70, 76.61)	29.15 (9.94, 42.43)
2 Years	10.87 (2.13, 33.71)	63.81 (41.31, 78.95)	25.31 (8.23, 38.08)
Subsample 2: 1914 – 2010			
0 Year	23.22 (7.00, 62.52)	66.28 (28.33, 80.72)	10.50 (0.55, 21.06)
1 Year	21.69 (6.67, 61.49)	67.14 (29.83, 80.31)	11.17 (0.88, 21.45)
2 Years	20.97 (6.68, 60.78)	67.37 (30.36, 80.10)	11.66 (1.09, 21.59)

Note: This model is estimated with 1 lag for both subsamples selected using BIC. Confidence intervals are obtained using the Kilian (1998) bias-corrected bootstrap method. The lower band indicates the 16th percentile of the bootstrapped distribution. The upper band indicates the 84th percentile.

TABLE 6

VARIANCE DECOMPOSITION RESULTS FOR *q*
(ROBUSTNESS MODEL 2 WITH NOMINAL INTEREST RATE DIFFERENTIAL)

Time Horizon	ε^s (%)	ε^d (%)	ε^m (%)
Subsample 1: 1798 – 1913			
0 Year	6.07 (0.65, 21.66)	64.37 (36.81, 85.66)	29.56 (4.71, 50.95)
1 Year	3.55 (1.09, 17.83)	67.60 (40.38, 87.41)	28.85 (5.01, 49.20)
2 Years	2.62 (1.56, 16.32)	68.75 (41.92, 87.24)	28.63 (5.38, 47.53)
5 Years	5.29 (3.58, 17.75)	73.13 (51.31, 86.32)	21.57 (5.01, 36.07)
10 Years	4.69 (3.84, 18.23)	77.77 (58.02, 87.28)	17.54 (4.44, 28.05)
Subsample 2: 1914 – 2010			
0 Year	1.02 (0.24, 9.77)	84.31 (54.82, 93.37)	14.67 (2.43, 39.33)
1 Year	1.95 (0.68, 12.75)	88.68 (61.87, 93.24)	9.37 (2.31, 30.25)
2 Years	1.53 (1.22, 12.63)	91.23 (66.08, 92.68)	7.24 (2.72, 25.04)
5 Years	3.21 (3.11, 16.64)	90.89 (66.47, 89.89)	5.90 (3.47, 20.84)
10 Years	11.19 (6.62, 27.79)	84.43 (61.23, 85.82)	4.38 (3.65, 15.26)

Note: The alternative specification is estimated with 6 lags. Confidence intervals are obtained using the Kilian (1998) bias-corrected bootstrap method. The lower band indicates the 16th percentile of the bootstrapped distribution. The upper band indicates the 84th percentile.

TABLE 7

VARIANCE DECOMPOSITION RESULTS FOR *q*
(ROBUSTNESS MODEL 3 WITH NOMINAL MONETARY BASE DIFFERENTIAL)

Time Horizon	ε^s (%)	ε^d (%)	ε^m (%)
Subsample 1: 1798 – 1913			
0 Year	2.61 (0.33, 14.09)	80.47 (65.26, 94.51)	16.93 (1.34, 25.48)
1 Year	1.27 (0.79, 13.43)	79.72 (64.66, 93.15)	19.00 (2.32, 25.85)
2 Years	1.18 (1.28, 13.98)	79.77 (64.96, 92.37)	19.05 (2.61, 24.32)
5 Years	5.21 (3.70, 25.30)	82.47 (64.58, 89.79)	12.32 (2.42, 15.07)
10 Years	4.26 (4.71, 25.22)	84.48 (66.21, 88.84)	11.26 (3.02, 12.15)
Subsample 2: 1914 – 2010			
0 Year	0.52 (0.36, 16.08)	98.65 (63.89, 96.46)	0.82 (0.43, 22.61)
1 Year	1.19 (0.80, 17.64)	96.18 (62.87, 95.01)	2.62 (1.05, 22.73)
2 Years	1.23 (1.11, 18.85)	96.19 (63.58, 94.58)	2.58 (1.26, 21.62)
5 Years	1.63 (2.43, 22.55)	95.48 (61.48, 91.54)	2.89 (2.27, 19.67)
10 Years	10.83 (8.34, 32.49)	84.68 (55.55, 83.51)	4.49 (3.44, 17.17)

Note: This model is estimated with 9 lags. Confidence intervals are obtained using the Kilian (1998) bias-corrected bootstrap method. The lower band indicates the 16th percentile of the bootstrapped distribution. The upper band indicates the 84th percentile.

TABLE 8

RESULTS FROM COUNTERFACTUAL EXPERIMENT

	Sample Standard Deviation		Standard Deviation from VAR Estimation		Standard Deviation Using Counterfactual Experiments	
	Subsample 1: 1795-1913	Subsample 2: 1914-2010	Subsample 1: 1795-1913	Subsample 2: 1914-2010	Subsample 1 Propagation: 1795-1913 Subsample 2 Shocks: 1914-2010	Subsample 2 Propagation: 1914-2010 Subsample 1 Shocks: 1795-1913
Change in q	0.0676	0.0753	0.0665	0.0831	0.0806	0.0685

TABLE 9

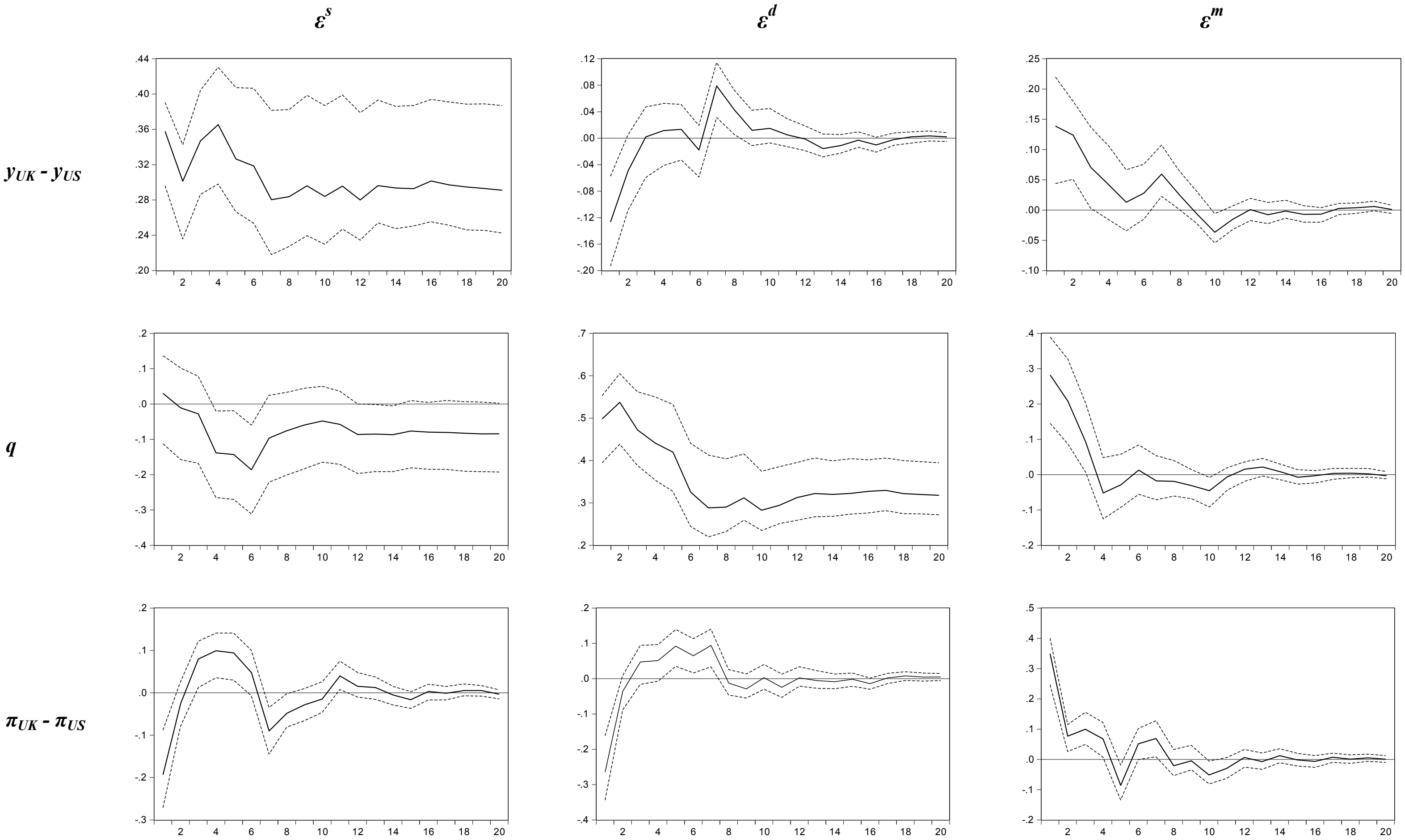
COUNTERFACTUAL VARIANCE DECOMPOSITION RESULTS FOR q
(BENCHMARK MODEL)

Time Horizon	ε^s (%)	ε^d (%)	ε^m (%)
Subsample 1 Propagation Subsample 2 Shock			
0 Year	1.88 (0.28)	88.53 (75.50)	9.59 (24.22)
1 Year	1.13 (0.16)	91.32 (81.23)	7.55 (18.62)
2 Years	0.86 (0.20)	92.99 (85.04)	6.15 (14.76)
5 Years	3.95 (5.25)	91.72 (85.39)	4.34 (9.36)
10 Years	4.14 (5.25)	92.33 (87.44)	3.53 (7.30)
Subsample 2 Propagation Subsample 1 Shock			
0 Year	0.47 (2.74)	88.27 (93.48)	11.26 (3.79)
1 Year	1.76 (4.56)	84.18 (90.75)	14.06 (4.69)
2 Years	1.53 (4.06)	86.89 (92.04)	11.57 (3.91)
5 Years	2.59 (3.59)	88.56 (93.25)	8.84 (3.15)
10 Years	11.66 (8.10)	80.62 (88.86)	7.72 (3.03)

Note: Original variance decomposition numbers for Subsample 1 and Subsample 2 from Table 4 reported in parentheses for comparison purposes.

FIGURE 1

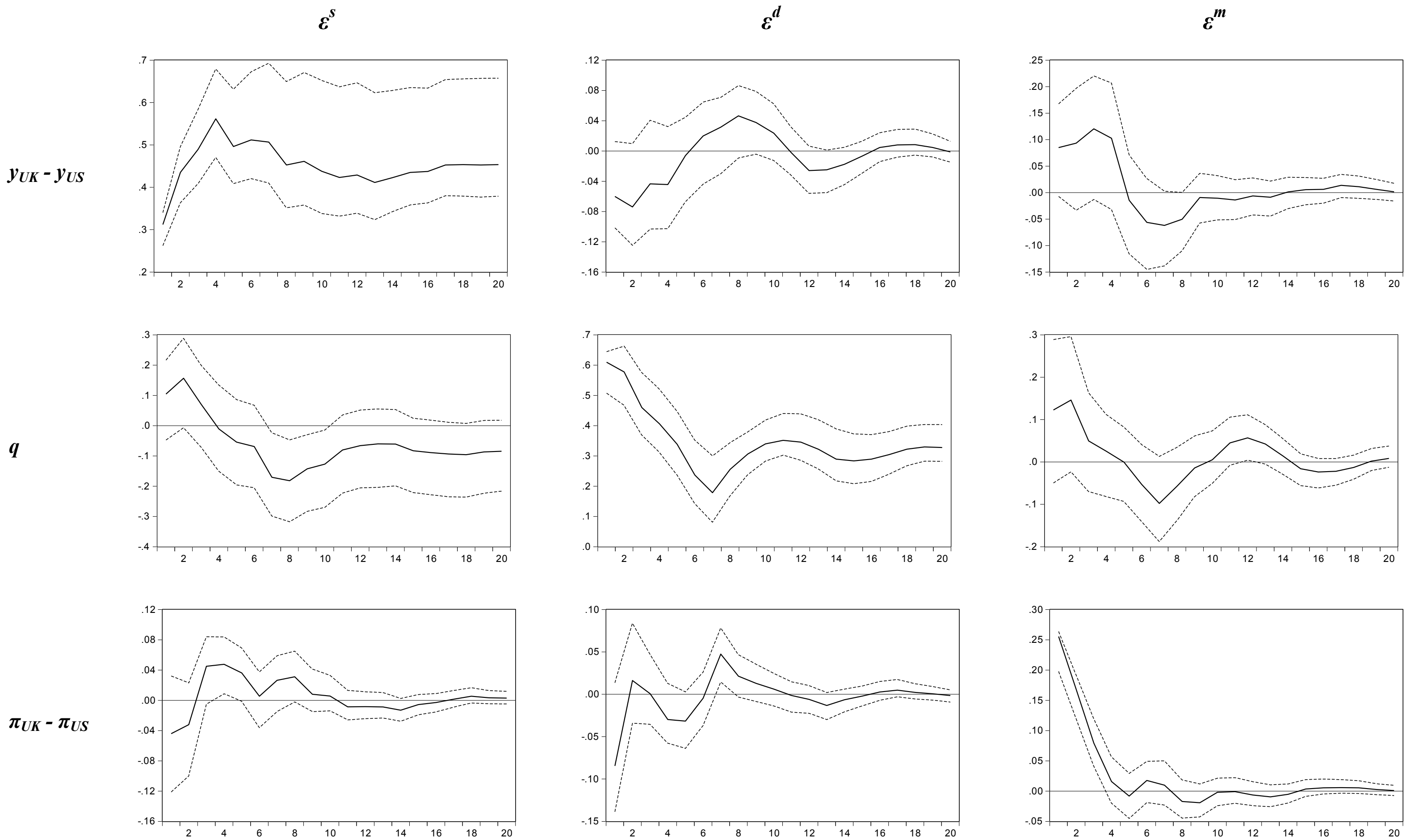
BENCHMARK MODEL IMPULSE RESPONSES (Subsample 1: 1795 – 1913)



Note: Benchmark model is estimated with 6 lags. The impulses are in response to a one standard deviation shock. Confidence intervals obtained using the Kilian (1998) bias-corrected bootstrap method. The lower/upper band indicates the 16th/84th percentile of the bootstrapped distribution.

FIGURE 2

BENCHMARK MODEL IMPULSE RESPONSES (Subsample 2: 1914 – 2010)



Note: Benchmark model is estimated with 6 lags. The impulses are in response to a one standard deviation shock. Confidence intervals obtained using the Kilian (1998) bias-corrected bootstrap method. The lower/upper band indicates the 16th/84th percentile of the bootstrapped distribution.

FIGURE 3

NORMALIZED STRUCTURAL NOMINAL SHOCK (ε^m) FROM BENCHMARK MODEL

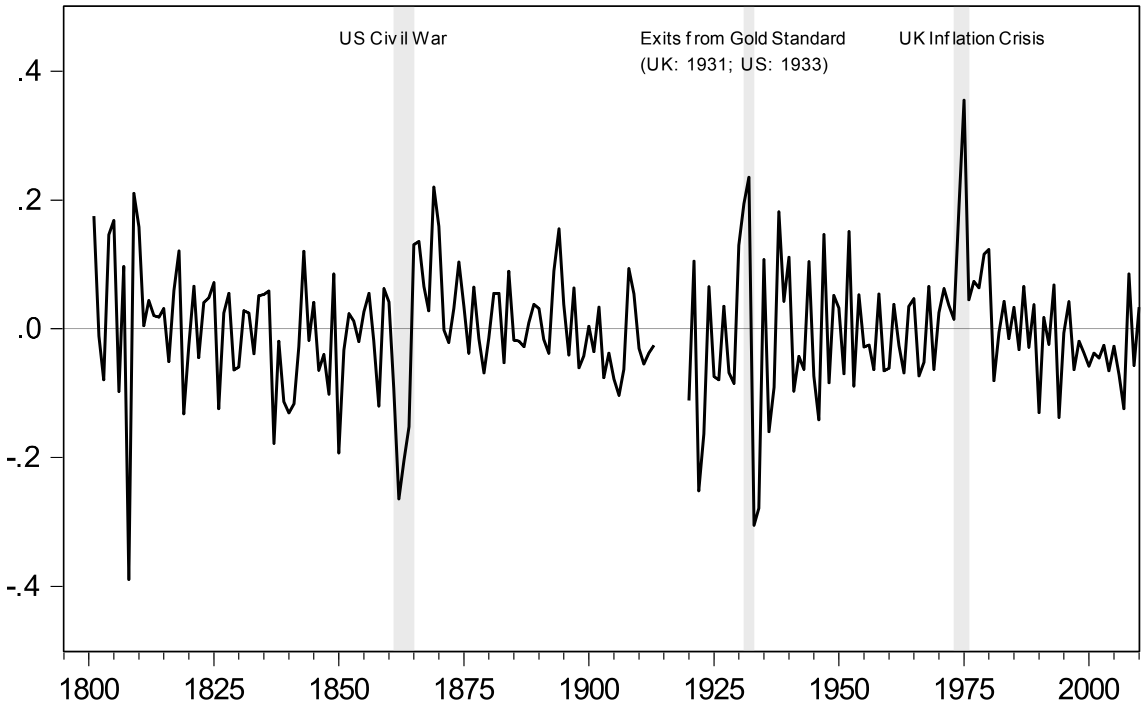
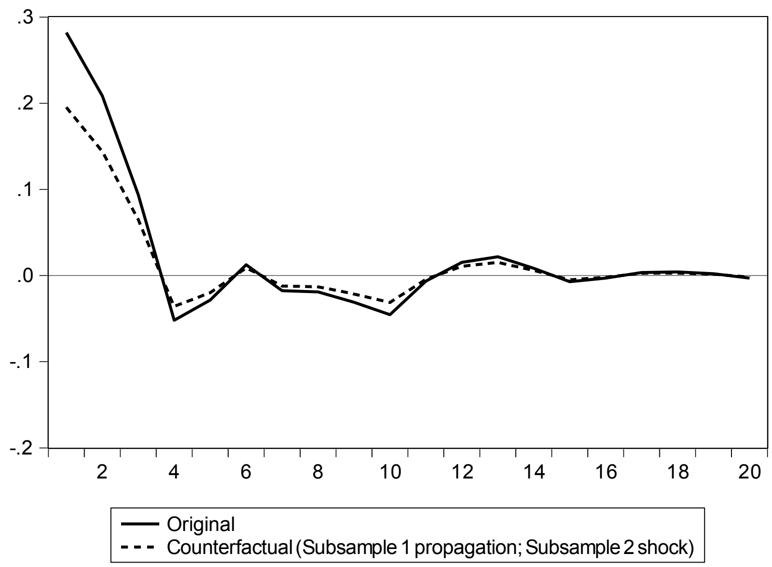


FIGURE 4

COUNTERFACTUAL IMPULSE RESPONSES OF *q* TO A ONE STANDARD DEVIATION NOMINAL SHOCK

Subsample 1 (1795-1913)



Subsample 2 (1914-2010)