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Globally correlated nominal fluctuations

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ABSTRACT

Fluctuations in nominal variables—aggregate price levels and nominal interest rates—are documented to be substantially more synchronized across countries at business cycle frequencies than fluctuations in real output. A transparent mechanism accounting for this striking feature of the nominal environment is described and quantitatively evaluated. It is based on the interaction between (small) cross-country spillovers of shocks, Taylor rules, and domestic no-arbitrage conditions. The mechanism is robust to various parameterizations and extensions aligning the model with other important aspects of domestic and international fluctuations. Furthermore, its key features are consistent with cross-country forecasts from Consensus survey.

1. Introduction

A growing empirical literature shows that domestic inflation rates contain a large 'global' component (Ciccarelli and Mojon, 2010; Mumtaz and Surico, 2012; Neely and Rapach, 2008). Put simply, in the post-WWII period or so, movements in inflation rates have been highly synchronized across countries. A large empirical literature has also established existence of the 'world business cycle'—recurrent fluctuations in real economic activity that tend to coincide across countries—and characterized its salient features.[1] Recently, a few studies have investigated how the two sides of global fluctuations, nominal and real, are statistically related to each other. Mumtaz et al. (2011) use a dynamic factor model to jointly identify a world cycle in both output growth and inflation. They find, among other things, that the 'world' factor is generally more important for movements in domestic inflation than in output growth. Wang and Wen (2007) document that cross-country correlations of inflation rates are much higher than cross-country correlations of cyclical movements in output, measured by deviations from HP-trend.

A part of the reason for the strong cross-country comovements of inflation rates uncovered by the literature is the similarity in inflation trends across industrialized economies, which occurred perhaps due to similar low-frequency changes in monetary policy: low inflation in the 1950s and 1960s, high in the 1970s, gradually declining in the 1980s, and low again since mid-1990s. Surprisingly, though, as documented in the first part of this paper, fluctuations in the nominal environment are substantially more synchronized across countries than fluctuations in real activity even at medium-term business-cycle frequencies; i.e., after low-frequency movements have been removed from the data. To the extent that

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[1] See, among others, Backus et al. (1992), Gregory et al. (1997), and Kose et al. (2003, 2008).

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at business cycle frequencies[2] domestic monetary policy is more likely to be able to control the domestic nominal environment than the real economy, it is not obvious why this should be the case. As surveyed below, existing quantitative-theoretical studies generally report relatively low cross-country correlations of nominal variables. Accounting for this feature of the data within a theoretical framework therefore seems important for our understanding of how domestic nominal variables are determined in an international environment—an issue that, in the face of accelerating globalization, has received considerable attention from policy makers.[3]

Our focus is on fluctuations in two key nominal variables, the aggregate price level and the short-term nominal interest rate—the monetary policy instrument. Our empirical observations are based on a sample of the largest industrial economies.[4] Using business cycle components—obtained with a band-pass filter—of aggregate price levels, short-term nominal interest rates, and real GDP, the fluctuations in the three variables are found to be similar in terms of their volatility and persistence, but markedly different in terms of their cross-country comovements.[5] Specifically, for the period 1960.Q1–2006.Q4 the average bilateral correlation of price levels is 0.52, that of short-term nominal interest rates 0.57, while that of real GDP is only 0.25. This empirical regularity is broadly robust to the inclusion of other economies as the required data become available, the exclusion of the Bretton Woods period of fixed exchange rates (ending in 1973), the exclusion of oil prices from price indexes, and to splitting the sample into two subsamples in 1984, the year generally associated with the start of greater focus by central banks on the stability of the domestic nominal environment. Furthermore, the findings are statistically significant.

The second part of the paper highlights a mechanism that can quantitatively account for this empirical regularity, yet appears to be consistent with a number of other important aspects of domestic and international aggregate fluctuations. The mechanism is based on cross-country spillovers of shocks, which are transmitted into nominal variables through an interaction between Taylor-type rules and no-arbitrage conditions for the returns on domestic capital (its marginal product) and domestic nominal bonds. This interaction generates a link between expected future states of the real economy and current nominal interest rates and prices: in equilibrium, the current interest rate and the price level depend on expected movements in domestic GDP and the return on domestic capital in all future periods, which due to cross-country spillovers of shocks over time are more aligned across countries than movements in current GDP. Corroborative evidence from outside of the model—based on the Consensus survey of professional forecasters—is provided, confirming such expectations.

The mechanism is investigated within a two-good, two-country business cycle model (Backus et al., 1994; Heathcote and Perri, 2002), where in each country a monetary authority follows a Taylor-type rule.[6] This is a useful benchmark for at least two reasons: (i) a large class of dynamic general equilibrium models in international macroeconomics builds on this framework, and (ii) the parsimonious nature of the model allows a transparent characterization of the mechanism. For our baseline experiments the model has only total factor productivity (TFP) shocks. Positive cross-country spillovers have been estimated for such shocks by Backus et al. (1992), Heathcote and Perri (2002), and Rabanal et al. (2011). In addition, Ahmed et al. (1993) and Crucini et al. (2011) provide empirical evidence that TFP shocks have been a dominant source of fluctuations in global economic activity. The choice of modeling monetary policy as a Taylor-type rule follows a large literature arguing that monetary policy of major central banks is reasonably well approximated by such a parsimonious feedback rule.[7]

Even though the mechanism requires positive spillovers of TFP shocks, it turns out that even some of the smallest estimates of such spillovers reported in the literature are sufficient to produce correlations similar to those in the data. The mechanism turns out to be quantitatively important also for a broad range of values of other parameters, the introduction of monetary policy shocks, and to whether the two-country TFP process is modeled as a stationary vector autoregressive (VAR) process or a cointegrated process with a vector error correction model (VECM) representation. In addition, when cross-country correlations of TFP shocks are chosen to match the dispersion of cross-country correlations for real GDP in the data, the model produces almost the same dispersion of cross-country correlations for the two nominal

[2] These are usually defined as frequencies with a periodicity between two and eight years.

[3] See, for instance, Bean (2006), Bernanke (2007), Mishkin (2007), and Besley (2008), among many others.

[4] Namely, Australia, Canada, Germany, Japan, the United Kingdom, and the United States for the period 1960.Q1–2006.Q4. In addition, from 1970.Q1 our sample includes also Austria and France and from 1976.Q1 Switzerland. Note that the sample is balanced in the sense that it is not biased towards European countries operating under fixed exchange rates within the European Monetary System (EMS) and, later on, the eurozone.

[5] Focusing on the price level allows us to remove low-frequency movements from the data with only one transformation. In contrast, filtering inflation leads to a double transformation of the price level, as inflation is already a transformation—first difference—of the data. In addition, it allows the same treatment of the price level as that of output, which in modern business cycle theory is usually filtered with either the Hodrick–Prescott or a band-pass filter to arrive at business-cycle frequencies. However, the empirical finding is not substantively changed when we use instead the inflation rate, the interest rate, and the GDP growth rate, or filtered inflation, interest, and GDP growth rates.

[6] Two studies have previously investigated cross-country correlations of nominal variables within two-country dynamic general equilibrium models. Kollmann (2001) constructs a model with sticky prices and wages, and technology and money supply shocks with cross-country spillovers. The model does not generate sufficiently high cross-country correlations for *both* prices and interest rates. A baseline version with only technology shocks and no nominal rigidities produces very low cross-country correlations for *both* nominal variables. Wang and Wen (2007) demonstrate that neither a prototypical sticky-price model nor a sticky-information model (both set off by money supply shocks) can generate, at the same time, sufficiently high cross-country correlations of inflation and realistic dynamics of inflation in relation to domestic output. A common feature of the two studies is that monetary policy is modeled as an exogenous stochastic process for money growth, rather than as a Taylor rule.

[7] See, among many others, Taylor (1993) and Clarida et al. (2000) for the United States, Clarida et al. (1998) for most of the G7 countries, and Nelson (2000) for the United Kingdom. Some studies estimate Taylor rules only for the post-1979 period, although others, for example Taylor (1999), Clarida et al. (2000), Nelson (2000), and Orphanides (2002), argue that a Taylor rule is a useful proxy for monetary policy also in the 1960s and 1970s, albeit with different parameter values than in the more recent periods.

variables as in the data. The mechanism is also robust to alternative assumptions about the structure of international asset markets. This is because *domestic* no-arbitrage conditions for real and nominal assets, rather than cross-country no-arbitrage conditions, are of first-order importance for the workings of the mechanism.

The baseline model, like many other models, falls short of accounting for some important features of international and domestic aggregate fluctuations. Two features that are of particular relevance, given our focus on nominal fluctuations, are: (i) the high volatility of nominal exchange rates in combination with smooth price levels, and (ii) the (strikingly similar across countries) correlations at various leads and lags of the domestic price level and the nominal interest rate with domestic output, which is henceforth referred to as the 'domestic nominal business cycle'. A number of recent studies[8] argue that exchange rate and interest rate dynamics over the business cycle are structurally related to cyclically moving distortions in Euler equations pricing financial assets, although the exact nature of such distortions (e.g., time-varying risk, limited participation) is still being debated. In an extension of the baseline economy such distortions are modeled as 'wedges', in the form of distortionary taxes, in the relevant Euler equations. Their joint stochastic processes with TFP are calibrated to match the observed exchange rate dynamics and the domestic nominal business cycle and the implications of such extensions for the cross-country correlations are investigated. The basic result is found to survive this robustness check.

The outline of the rest of the paper is as follows. Section 2 documents the empirical regularities, Section 3 introduces the baseline model, Section 4 presents quantitative findings for a baseline experiment, Section 5 explains the mechanism, Section 6 conducts thorough sensitivity analysis, and Section 7 concludes.

2. Properties of nominal business cycles

The empirical analysis is based on quarterly data for real GDP, consumer price indexes, and short-term nominal interest rates, usually yields on 3-month government bills, taken as a proxy for interest rates controlled by monetary policy. The data are for Australia, Canada, Germany, Japan, the United Kingdom, and the United States, for the period 1960.Q1–2006.Q4, plus Austria and France from 1970.Q1 and Switzerland from 1976.Q1.[9] With the exception of a robustness check regarding the detrending method, all statistics discussed in this section are for business-cycle components of the three variables obtained with the Christiano and Fitzgerald (2003) band-pass filter. Before applying the filter, the series for real GDP and the price levels were transformed by taking natural logarithms. Their fluctuations are thus measured as percentage deviations from 'trend'.

2.1. International nominal business cycles

We calculate bilateral cross-country correlations for the two nominal variables (i.e., the correlation of a country's variable with the same variable for each of the other countries) and compare them with those for real GDP. Fig. 1 provides a graphical representation of the main finding. It plots the bilateral correlations for the price level and the nominal interest rate against the bilateral correlations for output for the six-country sample going back to 1960.Q1. As is apparent, all but one point lie above the 45-degree line, meaning that for almost every country pair the bilateral correlations for the two nominal variables are higher than that for real GDP. Also note that the bilateral correlations for the two nominal variables are much less dispersed than those for real GDP.

The bilateral correlations for all country pairs are reported in Table 1 for the six-country sample and in Table 2 for the eight-country sample, which goes back to 1970.Q1 (a similar table for the nine-country sample going back to 1976.Q1 is omitted due to space constraints). In the six-country sample, for all 15 pairs the correlation of nominal interest rates is higher than that of output, and in all but one case the correlation of price levels is also higher. The mean (in the cross-section) bilateral correlations for the nominal interest rate and the price level are 0.57 and 0.52, respectively—about twice the mean bilateral correlation for real GDP, which is 0.27. In addition, the coefficients of variation (i.e., standard deviations divided by the corresponding mean) of the bilateral correlations for the nominal interest rate and the price level are 0.22 and 0.28, respectively, whereas the coefficient of variation of the bilateral correlations for real GDP is 0.89.

For each country pair Table 1 also reports (in parentheses) the 5th percentiles for $\phi_{ij}^R = \text{corr}(R_i,\ R_j) - \text{corr}(GDP_i,\ GDP_j)$ and $\phi_{ij}^p = \text{corr}(p_i,\ p_j) - \text{corr}(GDP_i,\ GDP_j)$, respectively. The percentiles are obtained by bootstrapping from the sample (Hardle et al., 2001) and provide a test of statistical significance that, for a given country pair, the observed bilateral correlations for the two nominal variables are higher than that for real GDP. A value of the 5th percentile greater than zero indicates that with 95% probability the 'true' correlation of nominal interest rates (or prices) is greater than that of real GDP. The percentiles are also computed for the mean values of the bilateral correlations in the cross-section. As can be seen, the bilateral correlations of nominal interest rates are significantly higher than those of output in 11 cases out of 15 and the

[8] Canzoneri et al. (2007), Atkeson and Kehoe (2008), Backus et al. (2010), and Sustek (2011).

[9] Accompanying supplementary material contains the data sources (http://www.sciencedirect.com/science/journal/0304393). For other developed economies, all required data are available only from either late 1970s or early 1980s. However, most of these additional economies are European economies that participated in the EMS and later adopted the euro. Including them into the sample would therefore bias the sample towards countries that operated under a fixed exchange rate for almost the entire sample period.

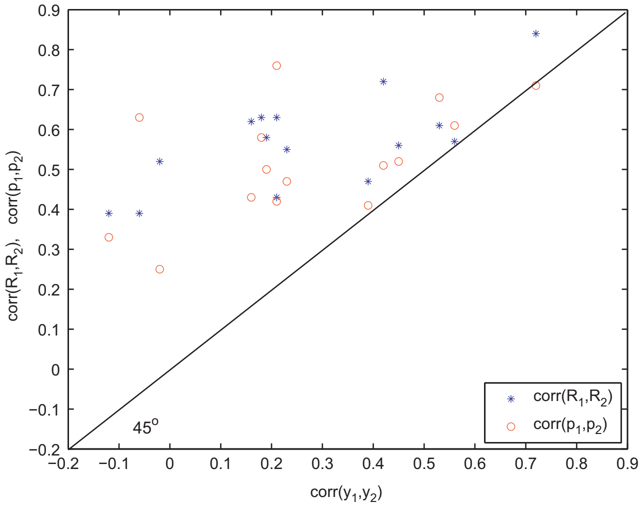

Fig. 1. Cross-country comovement of nominal variables vs. cross-country comovement of real GDP—the six-country sample (AUS, CAN, GER, JAP, UK, US), 1960.Q1–2006.Q4.

bilateral correlations of the price levels are higher in 10 cases. In addition, the cross-sectional mean bilateral correlations for both nominal variables are significantly higher than that for output.

These findings broadly hold also in the eight-country sample, as can be seen from Table 2. The mean bilateral correlations for the nominal interest rate and the price level are both 0.59, whereas that for real GDP is only 0.43 and these differences are statistically significant. The coefficients of variation are around 0.2 in the case of the two nominal variables, and slightly above 0.5 in the case of real GDP.

Even though the two nominal variables differ markedly from output in terms of the synchronicity of their cross-country comovements, they are comparably volatile and persistent. For example, in the six-country sample the mean standard deviations of output, the price level, and the nominal interest rate are, respectively, 1.39, 1.28, and 1.31[10] and the mean first-order autocorrelation coefficients are, respectively, 0.92, 0.94, and 0.91.

2.1.1. *Robustness checks: Bretton Woods, oil prices, policy change, and filtering*

For a part of the sample period—the Bretton Woods years—national monetary policies were constrained by governments' obligations to maintain fixed exchange rates with the dollar. It is well known that under fixed exchange rates the domestic economy is not insulated from nominal shocks originating abroad. In order to check that the high bilateral correlations for the price level and the nominal interest rate are not driven by the Bretton Woods period, Tables 1 and 2 report also the mean bilateral correlations and coefficients of variation for the period 1974.Q1–2006.Q4, which excludes the Bretton Woods years. However, as can be seen, the two summary statistics are little changed.[11]

Global commodity price shocks, in particular oil price shocks, may have been another source of the strong cross-country comovements of the price indexes. We therefore compute the bilateral correlations for CPI stripped off energy (and food) prices for those countries for which such data series are available for a long enough period. These are Austria, Canada, France, Germany, Japan, and the United States. The data, which for all six countries are available from 1970.Q1, largely come from Mumtaz and Surico (2012).[12] For 10 out of the 15 country pairs the correlations of prices are still higher than the correlations of real GDP. The mean bilateral correlation of the price levels is 0.6, while that of output is 0.5, and the difference is statistically significant. More importantly, the mean bilateral correlation of the price levels is little changed (it increases to 0.65) when the headline CPI is used for these countries (and this sample period) instead of CPI stripped off energy and food prices.

The previous robustness check addresses only the direct effect of energy prices on CPI. There may, however, be also indirect effects as oil prices feed into the prices of other CPI components as they work their way up through the supply chain. In order to address this issue, the cyclical component of a country's CPI is projected onto the cyclical component of Brent Crude oil price and then bilateral correlations for the orthogonal residuals obtained from these regressions are calculated. This analysis is carried out for the eight-country sample, 1970.Q1–2006.Q4, as in the 1960s oil prices were essentially flat. The mean bilateral correlation is found to decline mildly, from 0.59 to 0.58, and the coefficient of variation to

[10] The standard deviations of the nominal interest rates are for fluctuations measured in percentage points.

[11] Some researchers (Eichengreen, 1996) argue that during the Breton Woods period central banks retained a significant degree of monetary autonomy by imposing various capital controls, which reduced cross-border transmission of nominal shocks through fixed exchange rates.

[12] We thank Paolo Surico for providing us with the data. Note that this sample does not contain exactly the same countries as our benchmark six-country sample, but it is a subset of our eight-country sample.

Table 1
Cross-country correlations, 1960.Q1–2006.Q4.

	Aus	Can	Ger	Jap	UK
(a) Real GDP					
Can	0.53				
Ger	−0.02	0.16			
Jap	−0.12	−0.06	0.39		
UK	0.19	0.45	0.21	0.18	
US	0.23	0.72	0.42	0.21	0.56
Mean = 0.27 CV = 0.89					
Excluding Bretton Woods period Mean = 0.25 CV = 1.21					
Pre-1984 Mean = 0.34 CV = 0.71			Post-1984 Mean = 0.16 CV = 3.00		
(b) Nominal interest rates					
Can	0.61 (−0.03)				
Ger	0.52 (0.37)	0.62 (0.29)			
Jap	0.39 (0.35)	0.39 (0.30)	0.47 (−0.06)		
UK	0.58 (0.25)	0.56 (−0.02)	0.63 (0.28)	0.63 (0.30)	
US	0.55 (0.18)	0.84 (0.05)	0.72 (0.18)	0.43 (0.05)	0.57 (−0.11)
Mean = 0.57 (0.22) CV = 0.22					
Excluding Bretton Woods period Mean = 0.57 (0.23) CV = 0.26					
Pre-1984 Mean = 0.61 (0.16) CV = 0.20			Post-1984 Mean = 0.53 (0.27) CV = 0.37		
(c) Price levels					
Can	0.68 (0.04)				
Ger	0.25 (0.09)	0.43 (0.12)			
Jap	0.33 (0.26)	0.63 (0.55)	0.41 (−0.10)		
UK	0.50 (0.18)	0.52 (−0.06)	0.42 (0.08)	0.58 (0.26)	
US	0.47 (0.06)	0.71 (−0.12)	0.51 (−0.03)	0.76 (0.44)	0.61 (−0.06)
Mean = 0.52 (0.18) CV = 0.28					
Excluding Bretton Woods period Mean = 0.50 (0.16) CV = 0.29					
Pre-1984 Mean = 0.60 (0.18) CV = 0.17			Post-1984 Mean = 0.30 (0.03) CV = 1.03		

CV = coefficient of variation.
Note: the numbers in parentheses are the 5th percentiles for $\text{corr}(R_i, R_j) - \text{corr}(GDP_i, GDP_j)$ obtained by bootstraping.
Note: the numbers in parentheses are the 5th percentiles for $\text{corr}(p_i, p_j) - \text{corr}(GDP_i, GDP_j)$ obtained by bootstraping.

increase mildly, from 0.23 to 0.24, relative to the headline CPI data. When also four lags of the oil price are included in the regression, these statistics change to 0.53 and 0.3, respectively. In sum, the strong cross-country comovement of price levels documented in this paper is only marginally accounted for by fluctuations in the oil price.[13]

[13] Mumtaz and Surico (2012) work with inflation rates, rather than cyclical components of price levels, and arrive at essentially the same conclusion. Specifically, using a dynamic factor model, they find little correlation between oil price inflation and the 'global' component of national inflation rates.

Table 2
Cross-country correlations, 1970.Q1–2006.Q4.

	Aus	Aut	Can	Fra	Ger	Jap	UK
(a) Real GDP							
Aut	0.05						
Can	0.74	0.35					
Fra	0.21	0.70	0.53				
Ger	−0.10	0.68	0.18	0.47			
Jap	0.03	0.39	0.12	0.50	0.55		
UK	0.26	0.48	0.56	0.74	0.24	0.40	
US	0.39	0.56	0.69	0.67	0.53	0.52	0.65
Mean = 0.43 CV = 0.54							
Excluding Bretton Woods period							
Mean = 0.34 CV = 0.76							
Post-1984							
Mean = 0.19 CV = 2.07							
(b) Nominal interest rates							
Aut	0.63						
	(0.42)						
Can	0.65	0.64					
	(−0.18)	(0.14)					
Fra	0.48	0.57	0.49				
	(0.11)	(−0.29)	(−0.21)				
Ger	0.58	0.73	0.70	0.77			
	(0.48)	(−0.06)	(0.35)	(0.15)			
Jap	0.44	0.55	0.37	0.47	0.61		
	(0.26)	(0.01)	(0.08)	(−0.23)	(−0.08)		
UK	0.59	0.45	0.57	0.47	0.74	0.68	
	(0.16)	(−0.18)	(−0.12)	(−0.40)	(0.33)	(0.10)	
US	0.58	0.61	0.85	0.57	0.77	0.46	0.58
	(0.05)	(−0.10)	(0.07)	(−0.25)	(0.10)	(−0.23)	(−0.21)
Mean = 0.59 (0.07) CV = 0.20							
Excluding Bretton Woods period							
Mean = 0.55 (0.13) CV = 0.25							
Post-1984							
Mean = 0.46 (0.17) CV = 0.48							
(c) Price levels							
Aut	0.43						
	(0.20)						
Can	0.71	0.59					
	(−0.10)	(0.11)					
Fra	0.66	0.72	0.82				
	(0.34)	(−0.10)	(0.17)				
Ger	0.23	0.69	0.54	0.57			
	(0.10)	(−0.09)	(0.19)	(−0.07)			
Jap	0.36	0.57	0.62	0.78	0.52		
	(0.17)	(0.05)	(0.35)	(0.18)	(−0.16)		
UK	0.50	0.65	0.53	0.71	0.43	0.54	
	(0.10)	(0.04)	(−0.18)	(−0.13)	(−0.02)	(−0.03)	
US	0.44	0.69	0.67	0.76	0.59	0.77	0.57
	(−0.14)	(0.01)	(−0.15)	(0.01)	(−0.07)	(0.14)	(−0.22)
Mean = 0.59 (0.08) CV = 0.23							
Excluding Bretton Woods period							
Mean = 0.59 (0.18) CV = 0.24							
Post-1984							
Mean = 0.45 (0.16) CV = 0.56							

CV = coefficient of variation.
Note: the numbers in parentheses are the 5th percentiles for $\text{corr}(R_i, R_j) - \text{corr}(GDP_i, GDP_j)$ obtained by bootstraping.
Note: the numbers in parentheses are the 5th percentiles for $\text{corr}(p_i, p_j) - \text{corr}(GDP_i, GDP_j)$ obtained by bootstraping.

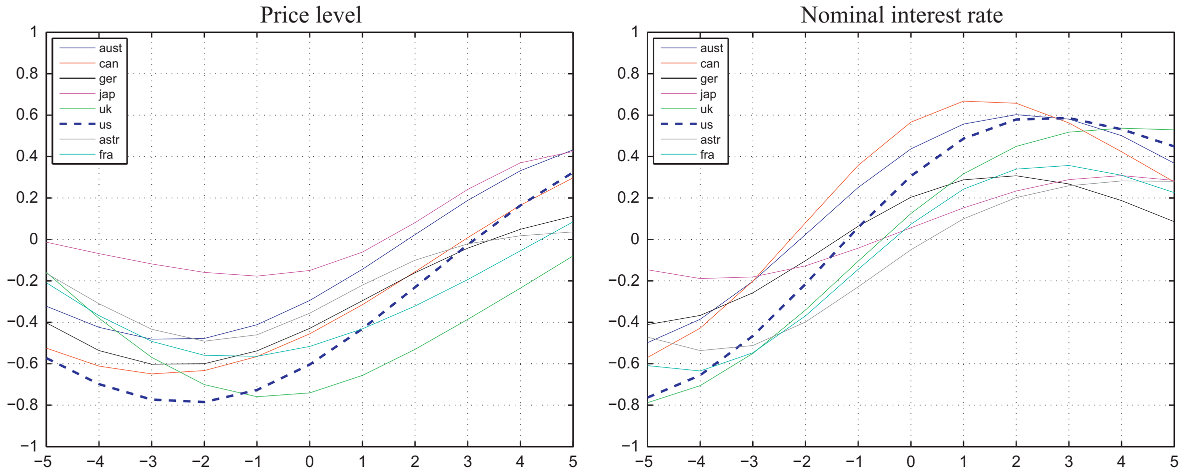

Fig. 2. Correlations of nominal variables in period $t+j$ with real GDP in period t.

In early 1980s, central banks in industrialized economies adopted a much tougher stance on inflation. In order to check if this policy change affects the cross-country correlations, the sample is split into two subsamples in 1984. (Note that the post-1984 period was also relatively free of large common commodity price shocks.) The summary statistics for the cross-section are contained in Tables 1 and 2. It is apparent that although the mean bilateral correlations declined after 1984 for all three variables, those for the two nominal variables remained substantially (and statistically significantly) higher than that for output. For example, in the eight-country sample, the post-1984 mean bilateral correlation for the nominal interest rate is 0.46, that for the price level is 0.45, while that for real GDP is only 0.19.

As explained above, our preferred way of measuring fluctuations in output and the price level is as band-pass filtered deviations from 'trend'. Our substantive finding is, however, robust to alternative detrending methods. For instance, in the case of the eight-country sample, when unfiltered nominal interest rates, CPI inflation rates, and GDP growth rates are used, the mean bilateral correlations are, respectively, 0.73, 0.56, and 0.15, with coefficients of variation equal to 0.16, 0.23, and 1.07. When the interest rates, inflation rates and GDP growth rates are all band-pass filtered, the mean bilateral correlations change to 0.59, 0.43, and 0.29, respectively, with coefficients of variation equal to 0.20, 0.35, and 0.77. In both cases the differences are statistically significant.

2.2. *Domestic nominal business cycles*

What are the properties of the two nominal variables in relation to domestic output? Kydland and Prescott (1990) have pointed out that a key characteristic of the nominal side of the U.S. business cycle is a countercyclical behavior of prices—i.e., the aggregate price level is negatively correlated with output over the business cycle. Backus and Kehoe (1992) and Mumtaz et al. (2011) extend this finding to a number of developed economies. The left-hand side panel of Fig. 2 confirms this regularity. The figure plots the correlation of a country's price level in period $t+j$ with its output in period t, for $j \in \{-5, -4, -3, -2, -1, 0, 1, 2, 3, 4, 5\}$. As is apparent, for all economies in the sample the contemporaneous correlation (i.e., that for $j=0$) is negative. Another striking feature of the data is, however, the systematic phase shift of the price level—in all countries the price level is more negatively correlated with future output than with current output; i.e., the price level tends to decline *ahead* of a GDP increase.

In the right-hand side panel of Fig. 2 this analysis is extended to the nominal interest rate. As in the case of the price level, the dynamics of the nominal interest rate exhibit a clear phase shift—the nominal interest rate is strongly negatively correlated with future output and positively correlated with past output. Although these dynamics of the nominal interest rate are well known for the United States (e.g., King and Watson, 1996), as in the case of the price level, it is striking that they extend to other developed economies too.

(*footnote continued*)
Besides 'cost-push' channels, oil price shocks can transmit into the price level also through 'general equilibrium' channels, which may be important in specific periods and can only be detected with a structural model. For instance, Backus and Crucini (2000) consider an international business cycle model in which oil is used in production. A positive oil price shock has a similar effect on output as a negative TFP shock in a typical model. In our model, TFP shocks transmit in equilibrium into the aggregate price level through monetary policy and asset market channels.

3. Baseline model

A world economy consists of two countries, denoted 1 and 2, which are populated by equal measures of identical, infinitely lived consumers. Producers in each country use country-specific capital and labor to produce a good, which will be referred to as a 'local' good. The good produced in country 1 is labeled by a, while that produced in country 2 is labeled by b. These are the only goods traded in the world economy. Within each country, goods a and b are combined to form a good that can be used for local consumption and investment, and which will be referred to as an 'expenditure' good.

3.1. Preferences and technology

Preferences of the representative consumer in country i are characterized by the utility function

$$\mathrm{E}_0 \sum_{t=0}^{\infty} \beta^t U(c_{it}, 1-n_{it}-s_{it}), \tag{1}$$

where $U(c, 1-n-s) = [c^{\mu}(1-n-s)^{1-\mu}]^{1-\gamma}/(1-\gamma)$, with $0<\mu<1$ and $\gamma \geq 0$, and where c_{it} is consumption, n_{it} is time spent working, and s_{it} is time spent in transaction-related activities. This 'shopping time' is given by the following parametric representation $s_{it} = \kappa_1[(p_{it}c_{it})/(m_{it})]^{\kappa_2}$, where $\kappa_1>0$ and $\kappa_2 \geq 1$, p_{it} is the domestic price level (i.e., the price of country i's expenditure good in terms of country i's money), and m_{it} is domestic nominal money balances.

Consumers supply labor and capital to domestically located, perfectly competitive producers, who have access to an aggregate Cobb–Douglas production function $z_{it}H(k_{it}, n_{it}) = z_{it}k_{it}^{\alpha}n_{it}^{1-\alpha} = y_{it}$. Here, z_{it} is a country-specific technology level, k_{it} is capital, y_{it} is output of the local good (either a or b), and $0<\alpha<1$. Technologies in the two countries follow a VAR(1) process $\lambda_{t+1} = A_0 + A\lambda_t + \varepsilon_{t+1}$, where $\lambda_t = [\ln z_{1t},\ \ln z_{2t}]^{\top}$ and $\varepsilon_{t+1} \sim N(0, \Sigma)$. Market clearing for goods a and b requires

$$a_{1t} + a_{2t} = y_{1t} \quad \text{and} \quad b_{1t} + b_{2t} = y_{2t}, \tag{2}$$

where a_{1t} is the amount of good a used by country 1, while a_{2t} is the amount used by country 2. Similarly, b_{1t} is the amount of good b used by country 1, while b_{2t} is the amount used by country 2.

Consumption and investment are composites of foreign and domestic goods

$$c_{1t} + x_{1t} = G(a_{1t}, b_{1t}) \quad \text{and} \quad c_{2t} + x_{2t} = G(b_{2t}, a_{2t}), \tag{3}$$

where x_{it} is investment, and $G(a,b) = (\omega_1 a^{-\rho} + \omega_2 b^{-\rho})^{-(1/\rho)}$, with $0<\omega_1<1$, $\omega_2 = 1-\omega_1$, and $\rho \geq -1$, is the Armington aggregator. Investment contributes to capital accumulation according to the law of motion $k_{i,t+1} = (1-\delta)k_{it} + x_{it}$, where $0<\delta<1$ is a depreciation rate. The prices of goods a and b in terms of the expenditure good of country 1 are determined competitively, and therefore given by the marginal products of these two goods

$$q_{1t}^a = \frac{\partial G(a_{1t}, b_{1t})}{\partial a_{1t}}, \quad q_{1t}^b = \frac{\partial G(a_{1t}, b_{1t})}{\partial b_{1t}}. \tag{4}$$

Similarly, the prices of the two goods in terms of country 2's expenditure good are given by

$$q_{2t}^a = \frac{\partial G(b_{2t}, a_{2t})}{\partial a_{2t}}, \quad q_{2t}^b = \frac{\partial G(b_{2t}, a_{2t})}{\partial b_{2t}}. \tag{5}$$

Using these prices, output of the two countries can be measured in terms of their respective expenditure goods as $q_{1t}^a y_{1t}$ and $q_{2t}^b y_{2t}$. It is easy to check that total expenditures in each country are related to this measure of output as $c_{1t} + x_{1t} + (q_{1t}^a a_{2t} - q_{1t}^b b_{1t}) = q_{1t}^a y_{1t}$ and $c_{2t} + x_{2t} + (q_{2t}^b b_{1t} - q_{2t}^a a_{2t}) = q_{2t}^b y_{2t}$, where the expressions in the parentheses are net exports. This is the model's measure of real GDP used by Heathcote and Perri (2002) and subsequent studies, which is also adopted here. Thus the following notation is used: $GDP_{1t} \equiv q_{1t}^a y_{1t}$ and $GDP_{2t} \equiv q_{2t}^b y_{2t}$.[14]

3.2. Monetary policy

A central bank in each country controls the nominal interest rate R_{it} on a one-period domestically traded bond, which pays one unit of country i's money in all states of the world in period t+1. The central bank sets the nominal interest rate according to a feedback rule

$$R_{it} = (1-\phi)[R + \nu_y(\ln GDP_{it} - \ln GDP) + \nu_{\pi}(\pi_{it} - \pi)] + \phi R_{i,t-1}, \tag{6}$$

where $\pi_{it} \equiv \ln p_{it} - \ln p_{i,t-1}$ is the inflation rate, and a variable's symbol without a time subscript represents the variable's steady-state value. In line with the literature, the policy rule features 'interest rate smoothing', captured by the parameter $0<\phi<1$. The central bank then supplies, through lump-sum transfers v_{it} to the consumer, the amount of nominal money balances dictated by the consumer's first-order condition for m_{it}. The nominal money stock thus evolves endogenously as $m_{it} = m_{i,t-1} + v_{it}$. We do not justify this monetary policy rule in terms of its welfare implications. It is simply taken as the

[14] The prices of the local goods are an order of magnitude less volatile in the model than y_t. As a result, the findings are not sensitive to whether GDP is measured in terms of the expenditure or the local good.

most parsimonious, empirically plausible, approximation of monetary policy in industrialized economies, as argued by the literature.

3.3. Consumer's budget constraint

Consumers hold money in order to economize on shopping time. In addition, they accumulate capital, a one-period internationally traded bond f_{it}, which pays one unit of good a in all states of the world in period t+1, and the domestically traded bond (described above), which is denoted by d_{it}.[15] Measured in terms of the domestic expenditure good, the consumer's budget constraint is

$$\frac{q_{it}^a f_{it}}{1+r_t^f} + \frac{d_{it}}{p_{it}(1+R_{it})} + \frac{m_{it}}{p_{it}} + c_{it} + x_{it} = q_{it}^{\varsigma}(r_{it}^k k_{it} + w_{it} n_{it}) + q_{it}^a f_{i,t-1} + \frac{d_{i,t-1}}{p_{it}} + \frac{m_{i,t-1}}{p_{it}} + \frac{v_{it}}{p_{it}}, \tag{7}$$

where r_t^f is the real rate of return (in terms of good a) on the internationally traded bond, r_{it}^k is the rate of return on domestic capital (equal in equilibrium to the marginal product of capital $z_{it}H_k(k_{it}, n_{it})$), w_{it} is the real wage rate (equal in equilibrium to the marginal product of labor $z_{it}H_n(k_{it}, n_{it})$), and q_{it}^{ς} is equal to q_{1t}^a in the case of country 1 and to q_{2t}^b in the case of country 2.

The relative purchasing power of consumers in country i is given by the real exchange rate. Following Heathcote and Perri (2002), the real exchange rate is defined as the price of the expenditure good of country 2 relative to the price of the expenditure good of country 1, i.e. $(1/q_{1t}^a)/(1/q_{2t}^a) = q_{2t}^a/q_{1t}^a$, which in equilibrium is also equal to q_{2t}^b/q_{1t}^b. An increase in this ratio represents an appreciation of the real exchange rate from country 1's perspective as less of this country's expenditure good (relative to the amount of country 2's expenditure good) is needed to purchase one unit of good a or b. The nominal exchange rate is consequently defined as

$$\chi_t = (q_{2t}^a/q_{1t}^a)(p_{2t}/p_{1t}). \tag{8}$$

3.4. Recursive competitive equilibrium

In each country, the consumer chooses state-contingent plans for c_{it}, $k_{i,t+1}$, m_{it}, d_{it}, f_{it}, and n_{it} in order to maximize (1) subject to (7), taking all prices as given. In all states of the world, the prices of capital and labor services, and of the two local goods a and b, are given by their respective marginal products derived above. In period t the state of the world economy is defined by the vector of technology levels λ_t, a vector of domestic endogenous state variables $\Upsilon_i = (p_{i,t-1}, R_{i,t-1}, k_{it}, \vartheta_{i,t-1}, f_{i,t-1})$, and a vector of foreign endogenous state variables $\Upsilon_j = (p_{j,t-1}, R_{j,t-1}, k_{jt}, \vartheta_{j,t-1}, f_{j,t-1})$, where $\vartheta_{i,t-1} \equiv d_{i,t-1} + m_{i,t-1}$, and similarly for $\vartheta_{j,t-1}$.

The equilibrium of the world economy is then characterized by a set of pricing functions for each country $\{r_i^k(\lambda, \Upsilon_1, \Upsilon_2)$, $w_i(\lambda, \Upsilon_1, \Upsilon_2)$, $q_i^a(\lambda, \Upsilon_1, \Upsilon_2)$, $q_i^b(\lambda, \Upsilon_1, \Upsilon_2)$, $p_i(\lambda, \Upsilon_1, \Upsilon_2)$, $R_i(\lambda, \Upsilon_1, \Upsilon_2)\}$, a set of aggregate decision rules for each country $\{n_i(\lambda, \Upsilon_1, \Upsilon_2)$, $k_i(\lambda, \Upsilon_1, \Upsilon_2)$, $m_i(\lambda, \Upsilon_1, \Upsilon_2)$, $d_i(\lambda, \Upsilon_1, \Upsilon_2)$, $f_i(\lambda, \Upsilon_1, \Upsilon_2)\}$, and a pricing function for the rate of return on the internationally traded bond $r^f(\lambda, \Upsilon_1, \Upsilon_2)$, such that the allocations and prices generated by these functions satisfy the consumer's optimization problem, the resource constraints (3), the goods-market-clearing conditions (2), a market-clearing condition for domestically traded bonds, $d_{it} = 0$, a market-clearing condition for the internationally traded bond $f_1 + f_2 = 0$, and the monetary policy rule (6).[16]

4. Baseline experiment

Table 3 reports the parameter values for the baseline experiment. Results of a thorough sensitivity analysis are reported in Section 6. As preferences and technology are the same as in Backus et al. (1994), the parameters of utility and production functions, and of the stochastic process for technology shocks, are either the same as in their paper or are calibrated to the same targets. The reader is therefore referred to their paper for details.

The parameters of the shopping time function are chosen to match moments of the empirical money demand function. In the model this function is given by the consumer's first-order condition for m_{it}, which in steady state has the form $(\kappa_1\kappa_2)(pc/m)^{\kappa_2}(p/m) = (1/w)[R/(1+R)]$. Setting κ_2 equal to 1 gives interest rate elasticity of m/p equal to −0.5, which is in line with empirical studies (Lucas, 2000), and setting κ_1 equal to 0.0054 implies annual velocity of money equal to 6.1—the

[15] As in Heathcote and Perri (2002), the denomination of the internationally traded bond has only second-order effects on equilibrium, which are not captured by our computational method. The denomination of the bond thus does not affect allocations and prices in the computed equilibrium. The model could be also extended to allow consumers in country i to hold the nominal bond of country j, but this would only clutter the model without affecting the computed equilibrium: the existence of such a bond would imply a no-arbitrage condition involving the nominal exchange rate, but this no-arbitrage condition (in its first-order approximation) holds even without such a complication, as shown below. Finally, as consumers of country i do not ‘shop’ in country j, they would never want to hold money of country j. Cross-border money holdings can thus be also abstracted from.

[16] Log-linear approximations to the equilibrium decision rules and pricing functions in the neighborhood of the model's non-stochastic steady-state are computed, using the method described by Hansen and Prescott (1995). Before computing the equilibrium, all nominal variables are transformed so that they are stationary. Following Heathcote and Perri (2002) a tiny quadratic cost of adjusting holdings of the internationally traded bond in the consumer's optimization problem is imposed in order to ensure stationarity of international bond holdings.

Table 3
Baseline calibration.

Symbol	Value	Definition
Preferences		
γ	2.0	Relative risk aversion
μ	0.34	Consumption share in utility
β	0.989	Discount factor
Technology		
δ	0.025	Depreciation rate
α	0.36	Capital share in production
ω_1	0.761	Weight on domestic good
ω_2	0.239	Weight on foreign good
$\sigma = 1/(1+\rho)$	1.5	Elasticity of substitution
Shopping time		
κ_1	0.0054	Level parameter
κ_2	1.0	Curvature parameter
Monetary policy rule		
π	0.0091	Steady-state inflation rate
ν_y	0.125	Weight on GDP
ν_π	1.5	Weight on inflation
ϕ	0.75	Smoothing coefficient
Process for technology shocks		
A_0	$(0.00072 \quad 0.00072)$	
A	$\begin{bmatrix} 0.906 & 0.088 \\ 0.088 & 0.906 \end{bmatrix}$	
Var $\varepsilon_1 = \mathrm{Var}\varepsilon_2$	0.00852^2	
$\mathrm{Corr}(\varepsilon_1, \varepsilon_2)$	0.258	

average U.S. annual velocity of M1 in the period 1959–2006. The estimates of the parameters of the monetary policy rule (6) vary greatly in the literature, depending on the countries considered, periods covered, and the exact specification of the rule. For the benchmark experiment, ν_π is set equal to 1.5 and ν_y is set equal to 0.125—the values used by Taylor (1993), which seem to be in the middle of the range of available estimates.[17] In addition, the steady-state inflation rate π is set equal to 0.0091—the average quarterly inflation rate in the United States between 1959 and 2006—and the smoothing coefficient ϕ equal to 0.75, a fairly common estimate.

Table 4, panel A, reports cross-country correlations of the price levels, the nominal interest rates, and real GDP in the model (entries labeled 'Baseline economy') and in the data (the cross-sectional averages). As in the case of the data, the artificial series generated by the model are filtered with the Christiano and Fitzgerald (2003) filter. The statistics for the model are averages for 100 runs of the model. It can be seen from the table that the model's quantitative predictions for the cross-country correlations are reasonably close to the data. In particular, in the model the cross-country correlations of the price levels and the nominal interest rates are 0.69 and 0.68, respectively, while the cross-country correlation of real GDP is only 0.23. In the data the mean values of these correlations are, respectively, 0.52, 0.57, and 0.27 for the six country sample, and 0.59, 0.59, and 0.43 for the eight-country sample.

Table 4, panel B, reports statistics for the domestic nominal business cycle and the nominal exchange rate.[18]. The model correctly generates a countercyclical price level and produces standard deviations of the price level and the nominal interest rate, relative to that of GDP, similar to those for the U.S. economy. However, the model fails to produce the empirical lead-lag pattern of the price level and the comovement between output and the nominal interest rate is far off the mark. As for the nominal exchange rate, it is clear that the exchange rate is substantially less volatile in the model than in the data and that its lead-lag relationship with real GDP is opposite to that in the data.[19] The failure of the model along these dimensions is not surprising—these are well known anomalies, at least for the United States, and we would therefore not expect the baseline model to account for them. In Section 6, however, the model is made consistent with these features of the data and it is asked if the model can still produce higher cross-country correlations of the two nominal variables than that of output.

[17] Taylor uses the weight on output equal to 0.5. This value is scaled down by four in the calibration in order to make it consistent with the nominal interest rate and inflation in the model, which are measured at quarterly rates, rather than annual rates.

[18] The business cycle properties of GDP, consumption, investment, net exports, terms of trade, and hours worked are similar to those reported by Backus et al. (1994) and Heathcote and Perri (2002) and are therefore omitted here (they are reported in Henriksen et al., 2009)

[19] The dynamics of real and nominal exchange rates in the data are essentially the same, as the ratio of the price levels in the data is relatively little volatile. A discussion of the observed dynamics of the nominal exchange rate is thus essentially equivalent to a discussion of the dynamics of the real exchange rate. In the model the volatility of the ratio of price levels is similar to that in the data. The low volatility of the nominal exchange rate in the model thus translates into low volatility of the real exchange rate.

Table 4
Findings for the baseline economy and its various extensions.

	Correlation		
	(p_1, p_2)	(R_1, R_2)	(GDP_1, GDP_2)
A. International correlations[a]			
Baseline economy	0.69	0.68	0.23
With monetary policy shocks	0.51	0.69	0.24
With domestic bond wedge	0.73	0.91	0.41
With foreign bond wedge	0.94	0.92	0.45
With foreign bond wedge and cap. adj. costs	0.65	0.30	0.19
Six-country sample, 1960.Q1–2006.Q4	0.52	0.57	0.27
Eight-country sample, 1970.Q1–2006.Q4	0.59	0.59	0.43

υ_{t+j}	Rel.	Correlations of GDP in period t with variable υ in period $t+j$								
	St. dev.[b]	$j=-4$	−3	−2	−1	0	1	2	3	4
B. Domestic correlations[a]										
Baseline economy										
Price level (p)	1.00	0.47	0.34	0.12	−0.13	−0.37	−0.52	−0.57	−0.54	−0.46
Nominal interest rate (R)	0.67	0.02	−0.24	−0.55	−0.82	−0.95	−0.88	−0.66	−0.37	−0.12
Nom. exchange rate (χ)	0.75	−0.42	−0.35	−0.21	−0.03	0.16	0.29	0.34	0.33	0.28
With domestic bond wedge										
Price level (p)	0.50	0.06	−0.17	−0.38	−0.43	−0.32	−0.05	0.25	0.41	0.41
Nominal interest rate (R)	0.85	−0.16	−0.29	−0.33	−0.19	0.10	0.44	0.67	0.68	0.48
With foreign bond wedge										
Nom. exchange rate[c] (χ)	1.86	−0.14	0.04	0.25	0.37	0.30	0.07	−0.17	−0.29	−0.24
Nom. exchange rate[d] (χ)	1.89	−0.28	−0.09	0.22	0.46	0.49	0.28	−0.03	−0.27	−0.32
U.S. economy										
Price level	0.82	−0.70	−0.77	−0.78	−0.73	−0.60	−0.43	−0.23	−0.03	0.16
Nominal interest rate	0.73	−0.66	−0.47	−0.22	0.06	0.31	0.49	0.58	0.59	0.53
Nom. exchange rate	3.09	0.23	0.21	0.21	0.20	0.19	0.17	0.14	0.08	0.01

[a] The entries for model economies are averages for 100 runs of the length of 188 periods each (each run is started off from a nonstochastic steady state, in which the price level is normalized to be equal to one, and the first 500 periods are discarded to limit dependence on the initial condition; the last 188 periods are kept). Except for the nominal interest rate, all artificial and data series are in logs; the nominal interest rate is expressed at annual rates. Before computing the statistics, the artificial and data series were filtered with the Christiano–Fitzgerald (2003) band-pass filter.

[b] Standard deviations are measured relative to that of GDP.

[c] No capital adjustment costs.

[d] With capital adjustment costs.

5. The mechanism

An understanding of the main result can be gained by plotting the responses of the model's variables to a 1% positive TFP shock. These responses are contained in Fig. 3. As the focus is on nominal variables, the responses of the real variables are described only briefly and the reader is referred to Backus et al. (1994) and Heathcote and Perri (2002). In this section it is helpful to abstract from the effects of nominal variables on the real economy. They occur due to an inflation tax, which affects the real money balances held by the consumer and thus the allocation of time between transactions, work, and leisure. These effects matter quantitatively only for ν_π sufficiently close to one (see Henriksen et al., 2009) and taking them into account at this stage would unnecessarily complicate the description of the mechanism without changing the main insight.

5.1. Real variables

Because the shocks in the two countries are correlated, a 1% increase in TFP in country 1 leads, on impact, to an increase in TFP in country 2 by 0.258%, where 0.258 is the correlation coefficient of the ε's. More importantly, due to spillovers, TFP in country 2 gradually catches up with TFP in country 1. As a result of a higher current and expected future TFP level, consumption in both countries increases, but it increases by less in country 2 than in country 1. There are two reasons for this. First, the net present value of country 2's future income is smaller than that of country 1. And second, there is intertemporal trade between the two countries: in order to take advantage of a higher TFP, country 1 increases investment by borrowing from country 2. This intertemporal trade is reflected in the decline of net exports of country 1, and the increase in the real return on the internationally traded bond. Because of the initially higher TFP in country 1, GDP is initially higher in country 1 than in country 2 (also labor is higher due to intertemporal substitution). As TFP in country 2 catches up with TFP in country 1, GDP in country 2 catches up with GDP in country 1. As a result of initially higher output in country 1, the price of good a falls, reflecting its abundance in the world market relative to good b. The terms of trade of country 1, defined as $e \equiv q_1^b/q_1^a = q_2^b/q_2^a$, therefore worsen, following the technology shock.

5.2. Nominal variables

The dynamics of the price level and the nominal interest rate are determined by the interaction between the Taylor rule and domestic no-arbitrage conditions. It is particularly convenient at this stage to work with log-linearized versions of the model's no-arbitrage conditions and the Taylor rule.[20] The Euler equations for capital and domestic and foreign bonds in country i are, respectively,

$$E_t\widehat{Q}_{it} + E_t\widehat{r}^k_{i,t+1} = 0, \tag{9}$$

$$\widehat{R}_{it} + E_t\widehat{Q}_{it} - E_t\widehat{\pi}_{i,t+1} = 0, \tag{10}$$

$$\widehat{r}^f_t + E_t\widehat{Q}_{it} + E_t\widehat{q}^a_{i,t+1} - \widehat{q}^a_{it} = 0, \tag{11}$$

where $\widehat{r}^k_{i,t+1} \equiv (r^k_{i,t+1} - r^k)/(1+r^k-\delta)$, $\widehat{R}_{it} \equiv (R_{it}-R)/(1+R)$, $\widehat{\pi}_{it} \equiv (\pi_{it}-\pi)/(1+\pi)$, and $\widehat{r}^f_t \equiv (r^f_t - r^f)/(1+r^f)$ are percentage deviations of the gross rates from steady state, and $\widehat{Q}_{it} \equiv \log Q_{it} - \log Q_i$ is the percentage deviation of the stochastic discount factor $Q_{it} \equiv \beta(U_{c,t+1} - U_{l,t+1}s_{c,t+1})/(U_{ct} - U_{lt}s_{ct})$. Combining Eqs. (9), (10), and (11), then gives, respectively, a no-arbitrage condition for domestic real and nominal assets and for real domestic and international assets

$$E_t\widehat{r}^k_{i,t+1} = \widehat{R}_{it} - E_t\widehat{\pi}_{i,t+1}, \tag{12}$$

$$E_t\widehat{r}^k_{i,t+1} = \widehat{r}^f_t + E_t\widehat{q}^a_{i,t+1} - \widehat{q}^a_{it}. \tag{13}$$

In addition, combining Eq. (13) for country 1 with that for country 2 gives a relationship between the return to capital in the two countries

$$E_t\widehat{r}^k_{1,t+1} + E_t(\widehat{q}^a_{2,t+1} - \widehat{q}^a_{1,t+1}) - (\widehat{q}^a_{2t} - \widehat{q}^a_{1t}) = E_t\widehat{r}^k_{2,t+1}, \tag{14}$$

where $E_t(\widehat{q}^a_{2,t+1} - \widehat{q}^a_{1,t+1}) - (\widehat{q}^a_{2t} - \widehat{q}^a_{1t})$ is the expected change in the real exchange rate. Notice that, due to the expected changes in the real exchange rate, the expected real returns to capital in the two countries do not need to be equalized. In addition, using a log-linearized version of Eq. (8) together with Eq. (12), Eq. (14) can be re-written in its nominal form as

$$\widehat{R}_{1t} + E_t\widehat{\chi}_{t+1} - \widehat{\chi}_t = \widehat{R}_{2t}. \tag{15}$$

This is, of course, the standard 'uncovered interest rate parity' condition.

[20] As preferences are standard time-additive, expected utility CRRA preferences and the exogenous processes are homoscedastic, little is lost by log-linearization. And in any case, the computed equilibrium in the quantitative experiments is for a log-linearized approximation of the model.

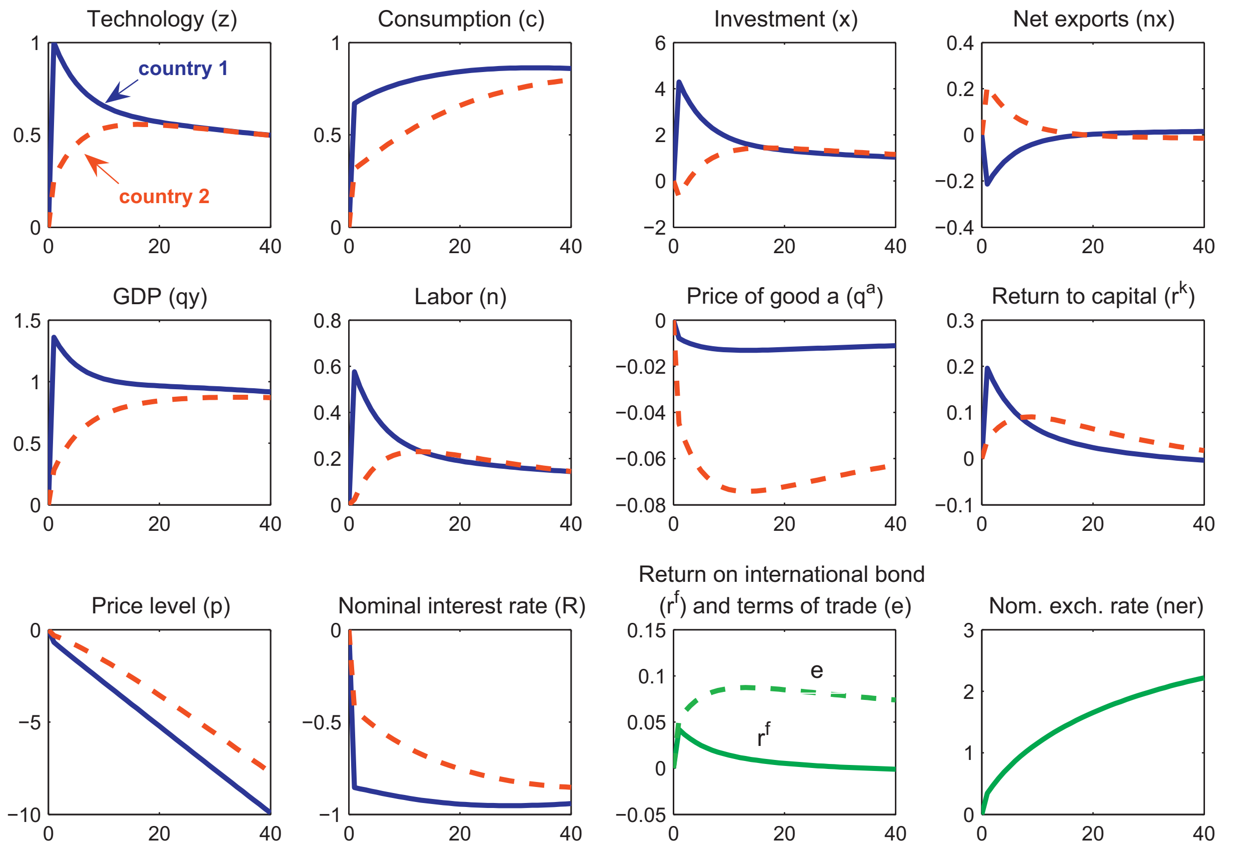

Fig. 3. Responses to a 1% technology shock in country 1 for the baseline calibration; rates of return are measured as percentage point deviations from steady state at annual rates; all other variables as percentage deviations.

5.2.1. Equilibrium prices and interest rates

As noted above, it is useful to abstract from the small inflation tax effects and think of real variables as being determined independently of nominal variables. Thus, given the equilibrium real quantities and relative prices, whose dynamics were described above, the equilibrium nominal interest rate and the price level of country i are determined by the no-arbitrage condition (12) and the (log-linearized) Taylor rule: $\widehat{R}_{it}=\tilde{\nu}_\pi\widehat{\pi}_{it}+\tilde{\nu}_y\widehat{Y}_{it}$, where $\tilde{\nu}_\pi \equiv \nu_\pi(1+\pi)/(1+R)$, $\tilde{\nu}_y \equiv \nu_y/(1+R)$, and $\widehat{Y}_{it} \equiv \log GDP_{it} - \log GDP$, and where for expositional purposes $\phi=0$ (the argument goes through even for non-zero ϕ). Combining the no-arbitrage condition (12) with the Taylor rule yields a first-order difference equation for inflation: $E_t\widehat{r}^k_{i,t+1}+E_t\widehat{\pi}_{i,t+1}=\tilde{\nu}_\pi\widehat{\pi}_{it}+\tilde{\nu}_y\widehat{Y}_{it}$. For $\nu_\pi>1$, this equation can be solved uniquely by forward substitution, excluding unstable equilibria that lead to either hyperinflations or hyperdeflations.[21] In a recursive equilibrium, $\widehat{p}_{i,t-1}$ is a state variable and $\widehat{p}_{it}$ is obtained as

$$\widehat{p}_{it}=\widehat{p}_{i,t-1}+E_t\sum_{j=1}^{\infty}\left(\frac{1}{\tilde{\nu}_\pi}\right)^j\widehat{r}^k_{i,t+j}-\tilde{\nu}_y E_t\sum_{j=1}^{\infty}\left(\frac{1}{\tilde{\nu}_\pi}\right)^j\widehat{Y}_{i,t+j-1}, \tag{16}$$

where the 'discount factor' in the summations is the inverse of the weight on inflation in the Taylor rule. The current price level thus reflects the expected future paths of output and the real return to capital. Substituting the price level from Eq. (16) into the log-linearized Taylor rule then yields a similar expression for the nominal interest rate

$$\widehat{R}_{it}=E_t\sum_{j=1}^{\infty}\left(\frac{1}{\tilde{\nu}_\pi}\right)^{j-1}\widehat{r}^k_{i,t+j}-\tilde{\nu}_y E_t\sum_{j=1}^{\infty}\left(\frac{1}{\tilde{\nu}_\pi}\right)^j\widehat{Y}_{i,t+j}. \tag{17}$$

An intuitive interpretation of Eq. (16) is as follows: period-t price level has to be consistent, according to the Taylor rule, with period-t output and the nominal interest rate, which (through the no-arbitrage condition (12)) has to be consistent with the expected real return to capital and expected inflation between periods t and $t+1$. Period-$(t+1)$ price level in turn has to be consistent with period-$(t+1)$ output and the nominal interest rate, which has to be consistent with the expected

[21] Hyperinflations in the model are costly because they make agents spend an increasingly larger amount of time in transaction-related activities, while hyperdeflations are costly because they lead to depletion of capital.

Table 5
Consensus survey: cross-sectional variance of real GDP growth forecasts.

Forecast horizon (years ahead)										
0	1	2	3	4	5	6	7	8	9	10
1.58	0.63	0.47	0.43	0.40	0.32	0.30	0.30	0.30	0.30	0.30

Note: the forecasts are for AUS, CAN, FRA, GER, JAP, UK, and US. The table reports the 1991–2010 average cross-sectional (i.e., across countries) variance of GDP growth forecasts for each forecast horizon.

real return to capital and expected inflation between periods t+1 and t+2, and so on. As a result of this recursion, period-t price level has to reflect all future states of the real economy.[22]

Notice that only the *domestic* no-arbitrage condition was used to arrive at Eqs. (16) and (17). The structure of international asset markets (i.e., the degree of their completeness), which affects the form of the cross-country no-arbitrage condition, affects domestic prices and the nominal interest rate only indirectly, to the extent to which it affects the dynamics of y_{it} and r_{it}^k.

With these derivations in hand, we can now go back to Fig. 3. As follows from Eq. (14), due to real exchange rate movements, cross-country borrowing and lending does not necessarily equate the returns to capital in the two economies. This is indeed the case in the benchmark experiment, as can be seen from the figure: the return to capital in country 1 increases on impact, while the return to capital in country 2 increases only gradually as TFP in country 2 catches up with TFP in country 1. The expected discounted sums of the rates of return in the two countries nevertheless increase on impact, as in both countries the return to capital is expected to stay above its steady-state level for much of the duration of the TFP shock. A similar argument also applies to the expected discounted sums of output. Thus, although output differs across the two countries between the impact period and the time when country 2 catches up with country 1, the discounted sums increase on impact in both countries. Because the price level and the nominal interest rate depend on the *difference* between the expected discounted sums of returns to capital and GDP, the sign of the responses of the two nominal variables depends on the relative weight on GDP in the Taylor rule. It turns out that, for the baseline calibration, the weight on GDP is sufficiently large to produce a fall in price levels and nominal interest rates in the two countries following the TFP shock in country 1. As a final observation, recall that the movements in the nominal interest rates have implications for the nominal exchange rate dynamics through the uncovered interest rate parity (15). As the nominal interest rate of country 1 is below that of country 2 in the plots in Fig. 3, the nominal exchange rate is increasing (i.e., appreciating from the perspective of country 1).

5.3. Cross-country GDP forecasts from Consensus survey

As the key element of the cross-country comovement of price levels and nominal interest rates in the model is expectations for similar paths of real GDP across countries in the longer run, it would add credibility to the mechanism if such expectations were observed outside of the model. Consensus Economics collects data on forecasts of professional forecasters for real GDP growth for a cross-section of countries. These forecasts are for the current year and every year up to a ten-year forecast horizon. The Consensus survey covers all countries from the eight-country sample, except Austria, from 1991 and the data are collected every six months. Table 5 reports the seven-country cross-sectional variance of these forecasts as a function of the forecast horizon (these are averages for 1991–2010). As can be seen, the variance declines with the forecast horizon, indicating that the forecasters expect similar growth across countries in the long run, despite expectations of different growth in the short run.

6. Sensitivity analysis

In order to check robustness of the mechanism, a thorough sensitivity analysis is carried out. Due to space constraints, only those experiments that lead to noticeable changes in the main quantitative result for the cross-country comovement reported in Section 4 are presented here. This is followed by findings for two extensions of the baseline model. The result of Section 4 is found to be only little sensitive to varying the elasticity of substitution between home and foreign goods (σ), the steady-state import share of GDP (b_1/y_1), and the shopping-time parameters (κ_1, κ_2). Also changes in the weight on inflation (ν_π) and the degree of interest rate smoothing (ϕ) in the Taylor rule have only small effects, except for values sufficiently close to the boundaries of permissible values, and the results for these experiments are reported in Henriksen et al. (2009).[23]

[22] The equilibrium price level is supported by nominal money balances required by consumers to achieve real money balances dictated by their first-order condition for m_{it} and supplied by lump-sum transfers υ_{it} from the central bank. As Cooley and Hansen (1995) show, such lump-sum transfers are, in an environment like ours, equivalent to open market operations.

[23] In addition, as should be expected from the discussion in the previous section, changes in the structure of international asset markets have only small effects. In particular, in the case of financial autarky—i.e., no international borrowing or lending—the cross-country correlations for output, the price level, and the nominal interest rate change to 0.43, 0.84, and 0.86, respectively.

6.1. Alternative parameterizations and specifications

We start by introducing monetary policy shocks. These are obtained for each country in the eight-country sample for the period 1984.Q1–2006.Q4 as residuals from Eq. (6), evaluated at the baseline parameter values. The mean bilateral correlation of these residuals is 0.13 and the mean standard deviation is 0.0026. The Taylor rules in the baseline model are thus appended with ς_{it}, where $[\varsigma_{1t},\ \varsigma_{2t}]\sim N(0,\ \Sigma_R)$, with the diagonal elements of Σ_R equal to 0.0026^2 and the off-diagonal elements equal to 0.13×0.0026^2. The shocks are assumed not to be autocorrelated, as the mean first-order autocorrelation in the data is 0.11, and are also assumed to be uncorrelated with TFP shocks. It is apparent from Table 4, panel A (entries labeled 'With monetary policy shocks') that including these shocks brings the correlation of price levels closer to the data, while leaving the other two correlations essentially unaffected.

Next, the left-hand side chart in panel A of Fig. 4 plots the cross-country correlations for output and the two nominal variables for alternative values of ν_y, which is varied between −0.05 and 0.25—a range that covers most of the estimates reported in the literature.[24] It is apparent that except for a small interval between 0.025 and 0.06, the cross-country correlations for the price level and the nominal interest rate are higher than that for real GDP.

Panel B of Fig. 4 plots the cross-country correlations for different values of the spillover term in the transition matrix A, and two values of ν_y, 0 and 0.125, the latter being the baseline value. Backus et al. (1992) estimate the spillover term to be 0.088—our baseline value—whereas Heathcote and Perri (2002) obtain a much lower estimate of 0.025. A_{12} is therefore varied between 0 and 0.1. In all experiments the diagonal elements of A are adjusted so that the highest eigenvalue is the same as in the baseline experiment, thus keeping persistence of the shocks constant. As can be seen, except for the case of no spillovers, the two nominal variables are correlated more strongly across countries than output. Crucially, the gap between the cross-country correlations for the two nominal variables and that for GDP opens up rapidly as we move away from the case of no spillovers. For Heathcote and Perri's relatively small spillover estimate of 0.025 the model generates cross-country correlations for, respectively, output, the nominal interest rate, and the price level, equal to 0.26, 0.71, and 0.61, for the case of $\nu_y=0$, and equal to 0.29, 0.45, and 0.48, for the case of $\nu_y=0.125$.

While a number of studies in the international business cycle literature adopt the stationary VAR specification for the joint TFP process, others—following Baxter and Crucini (1995)—argue that the process is better characterized as cointegrated of order C(1,1). Rabanal et al. (2011) obtain estimates for a VECM representation of the cointegrated process for the United States and the 'rest of the world'. Although the individual TFP processes are nonstationary, the joint process implies convergence to a common stochastic trend. The speed of convergence is estimated to be equal to 0.007 and plays a role similar to the spillover term $A_{12}=A_{21}$ in the stationary VAR process.[25] The results for this process are reported in the right-hand side chart in panel A of Fig. 4. Here, relative to the VAR specification, the correlations have declined for all three variables. Nevertheless, for most values of ν_Y, including the baseline value of 0.125, the nominal variables are still more strongly correlated across countries than output. The reason why the main finding broadly holds is that even though the spillover term is now much smaller, its effect is permanent.

So far we have considered only symmetric TFP processes. In panel C of Fig. 4, A_{12} is fixed at the baseline value of 0.088 while A_{21} varies (A_{22} is being adjusted so as to keep the largest eigenvalue of the matrix A the same). In this case the gap between the correlations for output and the two nominal variables stays approximately the same as A_{21} changes, even though the absolute values change. In particular, even when A_{21} is equal to zero, strong cross-country comovement of the two nominal variables is induced by $A_{12}=0.088$.[26]

Finally, panel D of Fig. 4 contains the results for varying the correlation of the innovations in the VAR process for TFP. The correlation is varied from −0.2 to 0.9, which allows to approximately match the range of cross-country correlations of real GDP reported in Fig. 1 (note that in the model there is nearly one-to-one correspondence between the correlation of the innovations and the correlation of output). A striking finding is that the model generates dispersion of the cross-country correlations for the two nominal variables similar to that in Fig. 1. A regression line through the data points in Fig. 1 yields a slope coefficient of 0.33 for both nominal variables, while a linear approximation to the curve for the correlations in panel D of Fig. 4 yields a slope of 0.36.[27] Also note that the prediction of the model that, other things being equal, a reduction in cross-country correlations of output leads to a less than proportional reduction in the nominal correlations is, at least qualitatively, consistent with the changes in the correlations in the post-1984 period noted in Table 2.

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[24] These are the values reported in the literature divided by four in order to make them consistent with the inflation and interest rates in the model, which are expressed at a quarterly rate.

[25] The process they estimate is

$$\begin{bmatrix} \Delta \ln z_{1t} \\ \Delta \ln z_{2t} \end{bmatrix} = \begin{bmatrix} \iota \\ \iota \end{bmatrix} + \begin{bmatrix} \kappa \\ -\kappa \end{bmatrix} (\ln z_{1,t-1} - \ln z_{2,t-2} - \ln \xi) + \begin{bmatrix} \epsilon_{1t} \\ \epsilon_{2t} \end{bmatrix},$$

where $\epsilon_{it}\sim(0,\sigma_i)$ are innovations whose correlation is estimated to be zero and the estimate of κ is equal to −0.007.

[26] We have also experimented with asymmetric Taylor rules, given the baseline calibration of the other parameters. In particular, ν_y for country 2 was varied, while ν_y for country 1 was kept at the baseline value of 0.125. The two nominal variables turned out to be still more strongly correlated across countries than real GDP as long as ν_y for country 2 was not too small, specifically not below 0.065.

[27] Although only cross-sectional implications of varying the correlation between the shocks are studied here, it is noted that variations in the correlation can occur also over time, leading to time-varying cross-country correlations for a given country pair. The 1970s oil price shocks are an example when the correlations increased.

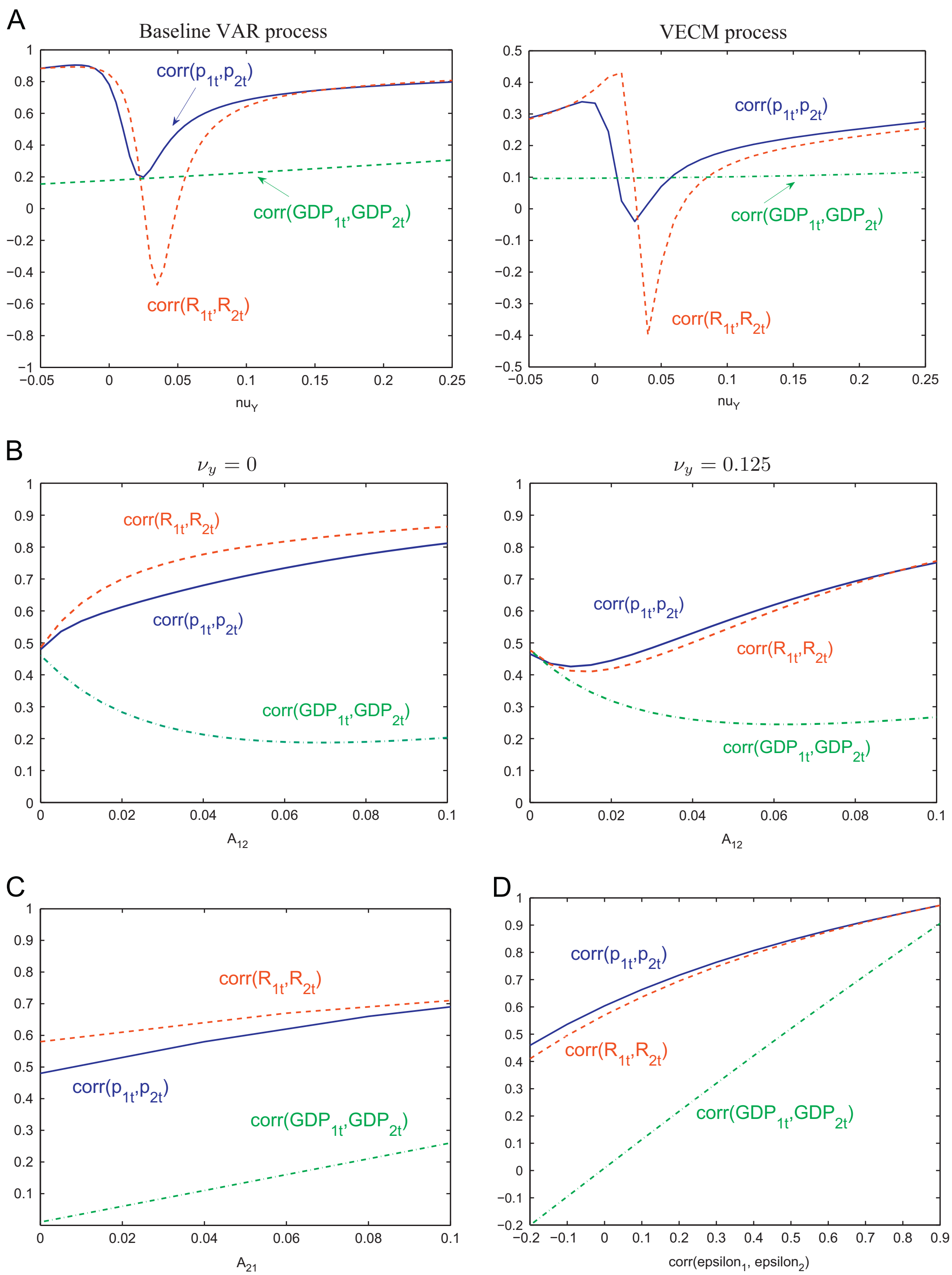

Fig. 4. Sensitivity analysis, baseline economy. The values of the baseline VAR process (panel A) are contained in Table 3; the values of the VECM are taken from Rabanal et al. (2011) and imply a tiny spillover equal to 0.007. (A) Weight on GDP in the Taylor rule (ν_y). (B) Cross-country spillovers of TFP shocks, VAR, symmetric ($A_{12} = A_{21}$). (C) Asymmetric spillovers ($A_{12} = 0.088$). (D) Correlation of TFP innovations.

6.2. Two extensions

As noted in the Introduction, recent research argues that the observed dynamics of interest rates and exchange rates are structurally related to cyclical distortions in standard Euler equations for financial assets. We are agnostic about the exact form of these distortions and model them as time-varying wedges. Due to space constraints, only an overview of these robustness checks is provided here and the reader is referred to Henriksen et al. (2009) for details.

First, a wedge in the Euler equation for domestic bonds is introduced into the model. It takes the form of an ad-valorem tax τ_{it}^{b} on investment in the domestic bond, d_{it}, with the tax revenues being rebated to the consumer in a lump-sum fashion.[28] A VAR(1) process for $[\ln z_{1,t+1},\ \tau_{1,t+1}^{b},\ \ln z_{2,t+1},\ \tau_{2,t+1}^{b}]$ is postulated and its parameter values are chosen by matching moments of the baseline VAR(1) process for $[\ln z_{1,t+1},\ \ln z_{2,t+1}]$—in particular, the autocorrelations of the two shocks and their mutual cross-correlations at various leads and lags—and the lead-lag pattern of the nominal interest rate for the United States reported in panel B of Table 4. The question asked is then if the model generates both the observed lead-lag pattern of the price level and higher cross-country correlations for the two nominal variables than that for output. The results are contained in Table 4, with entries labeled "With domestic bond wedge". They show that, similarly to the U.S. data, the price level now leads negatively and the model still produces higher cross-country correlations for the two nominal variables than that for output.

A similar experiment is carried out also for a wedge in the Euler equation for the internationally traded bond f_{it} (introduced into the model as an ad-valorem tax on investment in this bond, τ_{it}^{f}). It is important to note that as this wedge shows up only in the Euler equation for the foreign bond, the equilibrium price level and the nominal interest rate are still given by Eqs. (16) and (17), respectively. As a result, this wedge affects the price level and the nominal interest rate only indirectly by affecting output and the real return to capital (by affecting the real exchange rate and thus cross-country trade and borrowing and lending). A VAR(1) process for $[\ln z_{1,t+1},\ \tau_{1,t+1}^{f},\ \ln z_{2,t+1},\ \tau_{2,t+1}^{f}]$ is postulated and its parameters chosen to match the same moments of the baseline VAR process for TFP as in the previous case, the volatility of the nominal exchange rate (relative to that of real GDP), and the lead-lag pattern of the nominal exchange rate in relation to real GDP. The results are included in Table 4, under the label "With foreign bond wedge". By construction, the lead-lag pattern of the exchange rate is now more aligned with the data and the model accounts for a little over two thirds of exchange rate volatility observed in the data.[29] Even with this extension the model produces much higher cross-country correlations for the two nominal variables than that for output. The same experiment, but with capital adjustment costs $\varphi(k_{t+1}^{i}-k_{t}^{i})^{2}$, is also carried out in order to reduce the volatilities of investment and net exports, which increase above their respective volatilities in the data, once the foreign bond wedge is introduced into the model and the VAR process is calibrated as described above. The results of this experiments are reported in Table 4 under the label "With foreign bond wedge & cap. adj. costs" (this adds another moment, $\text{std}(x_t)/\text{std}(GDP_t)$, to be matched, in order to pin down φ). Again, the model produces higher cross-country correlations for the two nominal variables than that for output.[30]

7. Conclusion

There has been a growing interest in understanding the joint dynamics of real and nominal variables in the world economy. This study makes the empirical contribution of documenting that, at business cycle frequencies, fluctuations in aggregate price levels and nominal interest rates are substantially more synchronized across countries than fluctuations in output. This is an intriguing empirical regularity both from a theoretical point of view, as well as from the perspective of the policy debate about how the domestic nominal environment is determined in a globalized world. We then ask what is the key mechanism that brings about this striking feature of the data. To this end, an international business cycle model is employed that includes nominal variables and, in each country, a monetary authority that follows a rule with considerable empirical support. Even though the answer to our question very much involves dynamics, it is quite transparent within our abstraction. Due to spillovers over time of shocks across countries, expected future responses of national central banks to fluctuations in domestic output and inflation generate movements in current prices and interest rates that strongly co-move across countries even when output does not. International nominal business cycles are thus highly synchronized even when national monetary policies focus squarely on domestic output and inflation. A key element of our finding is that even a

[28] The Euler equation for d_{it} now becomes

$$E_t\{Q_{i,t+1}[(1+\tau_{i,t+1}^{b})/(1+\tau_{it}^{b})][(1+R_{it})/(1+\pi_{i,t+1})]\}=1$$

and the equilibrium price level (in its log-linearized form) becomes

$$\widehat{p}_{it}=\widehat{p}_{i,t-1}+E_t\sum_{j=1}^{\infty}(\tilde{\nu}_{\pi})^{-j}\widehat{r}_{i,t+j}-\tilde{\nu}_{Y}E_t\sum_{j=1}^{\infty}(\tilde{\nu}_{\pi})^{-j}\widehat{Y}_{i,t+j-1}+\Theta E_t\sum_{j=1}^{\infty}(\tilde{\nu}_{\pi})^{-j}\widehat{\tau}_{i,t+j-1}^{b}$$

where $\Theta>0$ is a function of the model's parameters.

[29] Although the nominal exchange rate is much more volatile now than it was in the baseline case, the ratio of the price levels is still fairly smooth (standard deviation of 2 for the former vs. 0.15 for the latter). Thus, as in the data, the volatilities of the nominal and real exchange rates in the model are about the same.

[30] Like many other models, the baseline model (as well as its two extensions) generally suffers from the Backus-Smith puzzle that, in contrast to the data, consumption in the model is more correlated across countries than output. A resolution of this puzzle usually involves various frictions. Henriksen et al. (2009) provide a discussion of how different types of frictions affect the mechanism highlighted in this paper. However, for some parameterizations (namely $A_{12}=A_{21}$ in the range of 0.005 and 0.015) the baseline model does produce the correct order of the cross-country correlations for output, consumption, and the two nominal variables.

modest degree of spillovers, in the range of the smaller estimates found in the literature, is sufficient to generate correlations such as those in the data.

Among the various extensions of the model one could possibly consider, one concerns the potential instability of the observed correlations. Backus and Crucini (2000) investigate the apparent instability in the correlation between the terms of trade and output. In a similar spirit, one could use our model, appropriately extended, to investigate the sources of instability in the correlations studied in this article and how they relate to the instability in other international correlations.

Having refrained from making the model more detailed than necessary for the question asked, while maintaining transparency of our answer, naturally there are some deviations in the data relative to the baseline model. The most notable ones, in our context, are in the forms of lead-lag properties of prices and nominal interest rates as well as exchange-rate volatility exceeding that implied by the model. Crucially, though, reduced-form extensions of the model suggest that these deviations can be substantially reduced without invalidating the findings. The importance of the mechanism should thus carry over to richer environments, in which explicit frictions have similar quantitative effects as our reduced-form wedges.

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Appendix A. Supplementary data

Supplementary data associated with this article can be found in the online version at http://dx.doi.org.10.1016/j.jmoneco.2013.05.006.

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