TSE working paper series

11-261

Research Group: *Macroeconomics*

December 2011

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World Technology Shocks and the Real Euro-Dollar Exchange Rate*

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December 15, 2011

Abstract

We extend the empirical SVAR literature on real exchange rates by extracting a common stochastic trend in productivity, interpreted as a permanent world technology shock. Overall, we find that innovations to world technology constitute an important, albeit not the dominant, source of movements in the real euro-dollar exchange rate. First, the dollar appreciates significantly in response to such an impulse. Second, the world technology shock accounts for approximately one-fifth of the variance of the forecast error in the real euro-dollar rate at business-cycle frequencies. Our results are in line with previous studies who find that demand or nominal shocks are the dominant sources of fluctuations in relative prices and provides limited support to productivity-based models of real exchange rate determination.

Keywords: Euro-Dollar Real Exchange Rate, World Technology Shocks, Structural VAR.

CEL Classification: C32, E32, F41

*All errors are mine.

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1 Introduction

How important are common world-wide technology shocks for fluctuations in real exchange rates? The aim of this paper is to address this issue in the context of the real euro-dollar rate, using a Structural Vector Autoregression (SVAR) approach with long-run restrictions.

The empirical literature on real exchange rates has long sought to identify those disturbances that are the primal source behind their movements. In a seminal paper, Clarida and Gali (1994) use long-run restrictions à la Blanchard and Quah (1989) to identify relative supply, demand and nominal shocks. They find that supply shocks - identified as the only disturbance that can permanently change the level of output in the long-run - have in general a very small contribution to the variance of the forecast error of real exchange rates at horizons between 1-20 quarters but shocks to (real) demand appear to be the most influential factor. Farrant and Peersman (2006) effectively repeat Clarida and Gali (1994) exercise using sign restrictions as their identification method. They conclude that the important role of demand shocks in explaining real exchange rate movements is significantly reduced and nominal shocks appear to be much more relevant. In line with Clarida and Gali (1994), they still find that supply shocks are the least influential factor; despite that their effect now appears to be significantly higher - for the EA-US relative price measure this ranges between 12-28% at horizons up to five years. Lewis (2006) also uses sign restrictions to identify *productivity* shocks, whose contribution to the variance of the forecast error in the real euro-dollar rate does not exceed 10%. On the other hand, Lastrapes (1992) finds an important contribution of "real" shocks to exchange rate movements, but his approach cannot distinguish between different types of real shocks. Further, Alquist and Chinn (2002), using cointegration techniques, and Schnatz et al (2003), using a general equilibrium model as a measuring device, suggest that productivity differences can be a significant source of variations in relative prices between US and Europe.

This paper extends the empirical SVAR-literature that seeks to identify sources of variations in real exchange rates. We use an SVAR model with long-run restrictions to identify a common stochastic trend in productivity, interpreted as a "permananent world technology shock". In this respect, our paper is new in many (related) ways. First, we emphasise on technology shocks, that is innovations that can permanently change the level of productivity, as opposed to shocks to the labour input as it is common in the literature[1]. Second, and most important, is our definition of a permanent technology shock: a common stochastic trend in productivity processes that constitutes the only source of unit-root. In other words, we assume, and test using statistical techniques, that productivity processes are one-to-one cointegrated. Rabanal et al (2008) show that the latter cannot be rejected at conventional confidence levels. Armed with this result, they use an estimated bivariate Total Factor Productivity (TFP) process for EA and US with a cointegrating vector $(1 -1)$ as the exogenous driving force of a standard, two-country DSGE model to show that they can match the observed real exchange rate volatility relative to output; when conventional International Business Cycle (IBC) models with stationary TFP processes do not (Backus et al 1994, Heathcote and Perri 2002). Bringing the common stochastic trend to productivity in the empirical SVAR literature that seeks factors of real exchange rate movements constitutes the main contribution of this paper. The importance of such a disturbance as a source of cyclical fluctuations in an open-economy context is also highlighted in Dupaigne and Fève (2009). They find that aggregate employment can increase in response to a world technology shock; thus casting doubts on Gali's (1999) famous result that employment decreases in the short-run after such an impulse. As they explain, this way of identifying *permanent* technology improvements - i.e. by using cointegrated productivities - is more immune to country-specific *stationary* disturbances and allows for open-economy dynamics. We use cointegrated productivities as a tool for identifying the "permanent world technology shock". To the extent that permanent common technology shocks are large in magnitude, and to the extend that technology is used differently across countries, they will have effects

[1] Note that shocks to the labour input can permanently change output and hours worked, but not labour productivity in the long-run. For a clear exposition see Gali (1999).

on macroeconomic variables at business cycle frequencies. Finally, our emphasis on a common permanent component in productivity meets a vast literature seeking to identify common factors in international cyclical movements (see for example Canova et al 2007, Kose et al 2003, Stock and Watson 2005)[2].

We estimate a SVAR model with US productivity growth, EA-US productivity differentials and the real euro-dollar exchange rate using annual data from the EU-KLEMS database from the Groningen Growth and Development Centre of the University of Groningen, for the period 1970-2007. We extract information about possibly several types of structural disturbances, but we are able to properly identify only one of them, namely a *permanent world technology shock*. Identification is achieved using the approach pioneered by Blanchard and Quah (1989) and employed by Shapiro and Watson (1988), Bayoumi and Eichengreen (1992), Ahmed et al (1993) and others. An important advantage of this approach is that we do not have to impose any restrictions on the dynamics. Rather, we use long-run restrictions which are in general less controversial (Ahmed et al 1993); even though their empirical validity has often been questioned (Farrant and Perrsman 2006).

From our analysis we derive two major conclusions. First, analysis of impulse responses suggests an important role of productivity for differences in relative prices across the Atlantic, even though the stochastic trend is common. Temporarily higher productivity in the US causes the dollar to appreciate significantly, and the effect is rather persistent. Our results are in line with Alquist and Chinn (2002) and Schnatz et al (2003) and provide further evidence for the existence of a Balassa-Samuelson type of effect in the euro-dollar rate, that might stem even from technology shocks that are common. This sustains the theoretical possibility that productivity improvements can induce wealth effects that are so large, that eventually outweigh supply-side effects and result in an increase in the relative price. Second, the world technology shock contributes more than 70% to the variance of the forecast error of US productivity

[2] Typically, factor models do not give any indications on what these factors are leaving one hungry for economic insights. However, Crucini et al (2008) find that productivity constitutes the main driving variable for the common component of the business cycle among the G-7; explaining around 47% of the common factor.

on impact, and less to the variability of differences in labour productivity across the Atlantic (around 25% at all horizons) but fails to be the dominant source of real-exchange rate fluctuations. Nevertheless, its contribution is still significant standing at 15.7% on impact and slightly more than one-fifth at business-cycle frequencies. Checks for robustness support this outcome, except for the case where the proxy for the real exchange rate is included in first-differences in our SVAR. These results do not contrast Clarida and Gali (1994) and others since the residual variability, around 80%, is explained by other factors, possibly demand-oriented, but disturbances that can permanently change productivity cannot be among them. The paper that is most similar to ours is Ahmed et al (1993) in so far that they identify and quantify the effects of a world supply shock (among other shocks) using long-run restrictions in a SVAR system for the US and a five-nation OECD aggregate. Since both output and employment are used in their SVAR system, this disturbance is similar to ours but we concentrate on the real euro-dollar rate. In comparison, the world supply shock is found to be a bit less significant; it explains around 10% of the variation in their measure of relative price but the contribution of country-specific shocks is a bit higher (16.9%). Their results are also in line to Clarida and Gali (1994) in so far that permanent demand shocks are dominant for variations in relative prices.

The next section describes the Blanchard and Quah (1989) methodology employed and our identifying restrictions. Section III discusses data and stationarity issues and Section IV provides the results. Some tests for robustness are done in Section V and section VI concludes.

2 Methodology

In this section we describe the Blanchard and Quah (1989) approach and how it can be fitted to our case. We impose - and confirm using statistical tests - that the log of US productivity (X_t^{US}) is non-stationary in levels but stationary in first differences, whereas the other two variables of our system - log of EA-US productivity differentials ($X_t^{EA} - X_t^{US}$) and the log of the real euro-dollar exchange rate (q_t) are stationary processes. We assume that there can be many disturbances in the economy, and our main identifying restriction that only one of them can have a permanent effect on the level of productivity allows us to extract information about the "world technology shock".

Letting $Y = (\Delta X_t^{US}, X_t^{EA} - X_t^{US}, q_t)'$ be the vector of endogenous variables and $\varepsilon = (\varepsilon_t^w, v_t^1, v_t^2)'$ be the vector of the world technology shock and two other unidentified shocks; then under standard regularity conditions each variable can be uniquely represented as an infinite distributed lag of these disturbances:

$$Y_t = A_0\varepsilon_t + A_1\varepsilon_{t-1} + A_2\varepsilon_{t-2} + ... \tag{1}$$

A_0 is the matrix of contemporaneous effects of the shocks on the endogenous variables, and thereafter $A_i \forall i \geq 1$ represent the effects at subsequent lags. The assumption that the shocks are uncorrelated implies that the variance-covariance matrix is diagonal, hence the assumption that $var(\varepsilon) = I_3$ is a simple normalisation. Our restriction that only the world technology shock can have a permanent effect on productivity implies that the second and third elements of the first rows of the sequence of A_i matrices sum up to zero:

$$\sum_{i=0}^{\infty} \alpha_{12i} = 0 \tag{2}$$

$$\sum_{i=0}^{\infty} \alpha_{13i} = 0 \tag{3}$$

For full identification, we need the long-run matrix to be lower-triangular. Thus, we need to impose a further restriction that the third element of the second row of the series of matrices A_i also sum to zero. Note that this restriction has no economic meaning since by stationarity of the differentials series; no shock can have a permanent effect.

$$\sum_{i=0}^{\infty} \alpha_{23i} = 0 \tag{4}$$

We show now how this structural representation can be retrieved from the data. Since Y follows a stationary process, it has a Vector Moving Average (VMA) representation:

$$
\begin{aligned}
&Y_t = e_t + B_1 e_{t-1} + B_2 e_{t-2} + ... \\
&var(e) = \Sigma
\end{aligned}
\tag{5}
$$

e_t represents the vector of canonical-form innovations and Σ the corresponding variance-covariance matrix. To convert equation (5) to the fundamental equation (1), we need to transform the innovations e_t into the structural shocks ε_t. Let the relation between the structural and the reduced-form innovations be linear, such that $e_t = C\varepsilon_t$ for some non-singular 3×3 matrix C. Comparing equations (1) and (5), we see that $A_0 = C$ and $A_1 = B_1 C$, $A_2 = B_2 C$ and so on, i.e. $A(L) = B(L)C$. Thus, knowledge of A_0 allows us to extract the structural shocks ε_t from the reduced-form innovations e_t and subsequently all A_j.

Is A_0 identified? A_0 has nine elements, thus we need nine restrictions to fully identify them. One can obtain six, using the relationship between the variance of the structural innovations (by assumption, the identity matrix) and the estimated reduced form variance as follows:

$$
var(e) = \Sigma \Rightarrow var(A_0 \varepsilon_t) = A_0 A_0^{'} = \Sigma \tag{6}
$$

The last three restrictions come from our identification restriction of a lower-triangular long-run matrix and hence from equations (2) - (4). This means that the matrix $(\sum_{i=0}^{\infty} B_i) A_0$ is lower triangular. Thus we can recover the structural system $A_1, A_2 ...$ and the world technology shock.

3 Data and Stationarity

Our sample consists of annual data from EU-KLEMS database from the Groningen Growth and Development Centre of the University of Groningen, for the period 1970-2007. Typically, SVAR-studies employ quarterly data but using data on annual frequency has its own advantages. Firstly, data on productivity *per hour* are not available (to our knowledge) at quarterly frequency for Europe and since productivity measured in units of hours worked controls better for labour intensity; we prefer to use this measure even at the cost of using a lower frequency sample. Second, since the workhorse of our model, i.e. that productivities are $(1 - 1)'$ cointegrated, is expected to hold true in the long-run or even at the steady state; lower frequency data might be better to look at. As Strauss (1996) puts it, simple "time disaggregation" from years to higher frequency is not likely to reveal long-run relationships. Finally, as Giannone et al (2008) point out, quarterly data for Europe are not very trustworthy and are only harmonised after 1991, questioning their use even at the gain of short-term dynamics.

We define EA as the eleven countries that first joined the monetary union, and are those countries used in the Are-Wide-Model (AWM) of the European Central Bank (ECB) developed by Fagan et al (2001). These countries are Belgium, Germany, Spain, France, Ireland, Italy, Luxembourg, Netherlands, Austria, Portugal and Finland. The main series used from the EU-KLEMS database were: "Gross Output at Current Basic Prices (millions of Euros)" code GO, "Total Hours Worked by Persons Engaged" code H-EMP, "Total Hours Worked by Employees" code H-EMPE, number of persons engaged and employees, and "Gross Output Price Indices" code GO-P. We also used from the "Total Database" of the Conference Board, obtained from the same source, population levels for each country. Data were not readily available to start directly estimations on them and had to be transformed. Details of these transformations are available on request.Our mainstream productivity series refers to the "persons engaged measure", using "persons employed" does not affect our results. Hours worked per employee and hours worked per person engaged both exhibit a downward trend, but hours should be a stationary series by definition since it is bounded by a physical constraint. To eliminate this problem we divided total hours by population. Our measure for US productivity is the "Output per hour" index of the Non-farm Business sector, seasonally adjusted, available from the Bureau of Labour Statistics. To estimate the real exchange rate we used values for the nominal rate of the "synthetic euro" from Datastream, whereas as a proxy of the price level we used data on consumer price indices. For the Euro Area, data come from the AWM of the ECB - transformed from quarterly to annual frequency by simple averaging[3]. For the US we used the annual consumer price index (US City Average, non-seasonally adjusted) available on the Bureau of Labour Statistics web-site. Note that the exchange rate is the value of one dollar in terms of the common European currency - thus an increase represents real euro depreciation. More details on the construction of the euro-dollar real exchange rate are found in Appendix A. All variables are presented and analysed in log-levels, and hence first differences represent annual growth-rates. Graphs of all the series are presented in Appendix A.

According to our specification, productivity in EA and US should be non-stationary integrated one processes who are mutually cointegrated. Graphs of the natural logarithm of both series are presented in Appendix A and both series exhibit an upward trend (Figure 8). Further, we observe that since 1970s Europe has been constantly more productive than the US, thus decreasing the productivity gap between the two regions. Moreover, the graph is suggestive of the existence of a common stochastic trend in the two series, similarly to Rabanal et al (2008). Also, Giannone et al (2008) provide evidence of a common trend between EA and US by showing that the gap between US and European GDP per capita has been more or less constant since 1970; indicative of some long-run relation in the two series[4]. In what follows, we show that the assertion of a common stochastic trend can be also supported on statistical grounds in our sample.

For each series we conducted three different types of stationarity tests: the Aug-

[3]Data from the AWM are up to 2005, thus the two missing years were obtained by extrapolating the sample using annual growth rates of the HCIP index from the ECB web-site.

[4]Figures 6 and 7 in Giannone et al (2008) depict that constant gap pretty clearly.

mented Dickey-Fuller test (ADF), the Phillips-Perron test (PP) and the Kwiatkowski-Phillips-Schmidt-Shin test (KPSS) so that we can test both the null hypothesis of a unit root and that of stationarity. The results of these tests are presented in the Appendix B. First, we check whether the log-productivities are integrated of order one - $I(1)$. For the EA, the null hypothesis of a unit root cannot be rejected at conventional levels by the ADF test and the KPSS test rejects the null of stationarity at 1% level in favour of a random walk specification. On the other hand, the PP test rejects the null hypothesis of a unit-root (p-value = 0.0073) against the alternative of stationarity. The results for the corresponding US series provide even more solid evidence for the existence of a unit root. The above, together with strong rejection of the unit-root once the series are differenced (not reported) support that the series are indeed $I(1)$. Therefore, we can now move on to statistical tests of cointegration between the two series.

According to Hamilton (1994), when the cointegrating vector is known the by far best method of checking for cointegration is to construct a series using the implied cointegrating relationship and check that series for stationarity. Following this logic, unit-root tests provide rather solid evidence for our intuitive belief that US and EA exhibit a common trend (see table 5 in Appendix B). The ADF and PP tests reject the null hypothesis of a unit root at 5% level in favour of the alternative, i.e. stationarity, with p-values equal to 2.6% and 2.2% respectively. Even though the KPSS is less supportive of these results, rejecting stationarity at 5% in favour of a random-walk specification, it does not do so at 1% level.

The problems about the intuitive belief of $(1-1)'$ cointegration come from the Johansen tests. In particular, with one lag in fist differences in the Vector Error Correction (VEC) model, and using constants in both the cointegrating relation and the data we cannot reject the null hypothesis of zero cointegrating relationships at all levels, with both Trace and Max-Eigenvalue tests; but the null that the vector is equal to $(1-1)'$ cannot be rejected. Introducing a dummy for the German unification significantly lowers the p-value for the null hypothesis of a zero cointegrating vector very close to 10%; while keeping the p-values of one cointegrating vector equal to

$(1 - 1)'$ at high levels, thus improving our prior belief. Moreover, removing the constant from the data fixes the results. In that case, both Trace and Eigenvalue Tests reject the null hypothesis of zero cointegrating vectors and cannot reject the null of one at 5% level; therefore supporting the existence of a single cointegrating vector (see Appendix B, Table 7). Further, the null hypothesis that that vector is $\beta = (1 - 1)'$ cannot be rejected with a high p-value (0.51). The latter, combined the outcomes of unit-root tests, the results of other studies that render productivity differentials stationary and last but not least our intuitive belief allow us to move on with the estimation of the SVAR. Before doing so, we check the stationarity of the real exchange rate series.

Whether the real exchange rate is a stationary process or a unit-root process is not a conclusive issue in the empirical literature; and to some extent neither in the theoretical one. Intuitively, the real exchange rate should be stationary by virtue of the Purchasing Power Parity (PPP). Consequently, any test checking for unit roots in real exchange rates is equivalently a test for the empirical validity of the PPP; which it is known to be a rather contentious issue. Froot and Rogoff (1994) analyse a number of studies that empirically test real exchange rates as unit-root processes and conclude that they appear to be stationary for as long as the horizon is sufficiently long. Second, it seems that the null hypothesis of a unit-root is more difficult to reject in the post Bretton Woods period of floating - and rather volatile - exchange rates rather than under fixed exchange rate regimes[5]. Third, annual data seem to favour more the stationarity of the real exchange rate (Strauss 1996). In our model we impose PPP in the long-run and that no shock can have permanent effect on the real exchange rate; with official tests giving solid evidence to the latter. Even though PP test confronts stationarity; the other two tests support it. Results of statistical tests are presented in Table 6 in Appendix B. In particular, the null hypothesis of a unit root is rejected by the ADF test at 10%, with a p-value at a bit lower than 7%, and KPSS strongly supports stationarity since the null cannot be rejected at any

[5] Lothian and Taylor (1994) when using only post Bretton-Woods data cannot reject the null hypothesis of a random walk for any real exchange rate series used, but that is "easily rejected" when the entire sample is used (more than 100 years).

level.

However, careful inspection of the Figure 8 in Appendix A reveals two-periods of large-scale dollar appreciation in the mid-80s and towards the end of the century which, at least for the latest episode, have been largely attributed to higher productivity growth in US than in EA (Alquist and Chinn 2002). Additionally, one should not forget that there can also be theoretical reasons why deviations from PPP might be long and persistent, with the most prominent one being the renowned Balassa-Samuelson effect (attributed to Balassa 1964 and Samuelson 1964). If productivity differences play an important role in determining relative prices thus causing persistent deviations from PPP, conventional unit-root tests can face real difficulty in distinguishing between a stationary but persistence process and a unit-root process. Problems of low power in these tests are well known. Indeed, Alquist and Chinn (2002) and Schnatz et al (2003) provide solid evidence for the existence of a Balassa-Samuelson effect in the real euro-dollar rate. Economy-wide productivity differentials "appear to have a strong impact on the dollar/euro rate", where "each percentage point in the US-Euro Area productivity differential results in a five percentage point appreciation of the dollar"[6]. The above call for the use of a slightly different type of test, the so called Covariate Augmented Dickey-Fuller (CADF) test developed by Hansen (1995). This test departs from the standard univariate context of unit-root tests (like ADF) and makes use of more variables that can potentially explain movements in the variable of interest. Hansen (1995) shows that including the appropriate covariates can result in enormous gains in power. Obviously, if productivity differentials is an important determinant of real exchange rate movements, and since we have previously established its stationarity, the series makes a prime candidate to be used as an additional explanatory variable in CADF tests. Performing the test we find that the null hypothesis of a unit-root in the (log-) real exchange rate is rejected at 5% level when we include the current and past value of EA-US productivity differentials in the regression, and at 1% level once we also include a future value (see Appendix B, Table 6). Hansen (1995) stresses the importance that the included covariate(s) is

[6] Alquist and Chinn (2002). Schnatz et al (2003) find this number to be lower at 1.5%-2%.

stationary, hence we also use US productivity growth to perform the test, our conclusions are unaltered. Thus, we can be confident that the real exchange rate - or at least the measure we have in hand - is a stationary, albeit persistent, process and include it in log-levels in our estimations.

Overall, stationarity and cointegration results do allow us to construct a SVAR system as outlined in the previous section. In what follows we provide the main results of this analysis.

4 Results from SVAR

In this section we analyse the Impulse Response Functions (IRFs) and the variance decompositions of the three variables in our system. We estimated a Vector Autoregressive (VAR) model with p-lags in the vector of the endogenous variables and we set $p = 1$ as chosen by both Akaike and Scwartz Information Criteria. Maximum Eigenvalue tests cannot reject the null hypothesis of zero cointegrating vectors between the endogenous variables.

As explained above, our identification restriction asserts that only the world technology shock can permanently change US productivity. Figure 1 presents the IRF of the level of US productivity to this disturbance. US productivity increases in response to a permanent innovation in world technology and follows a small hump shaped behaviour: it increases by slightly less than 1.2% on impact, reaches a peak after seven years (1.5%) only to decrease thereafter and restore itself to its new long-run value at a level 1.2% higher than before the shock. Figure 2 provides the IRFs of the other two variables to the permanent common technology shock. The world technology shock is primarily felt in the American economy, since the differentials series decreases on impact by 0.8% before it reverts back to its long run (constant) level.

Finally, we observe that temporarily higher productivity in the US causes a real euro depreciation. Alquist and Chinn (2002) construct an empirical model using cointegration between the euro-dollar real exchange rate and productivity differentials growth with quarterly data over 1986-2001, to find that a one percentage point in the

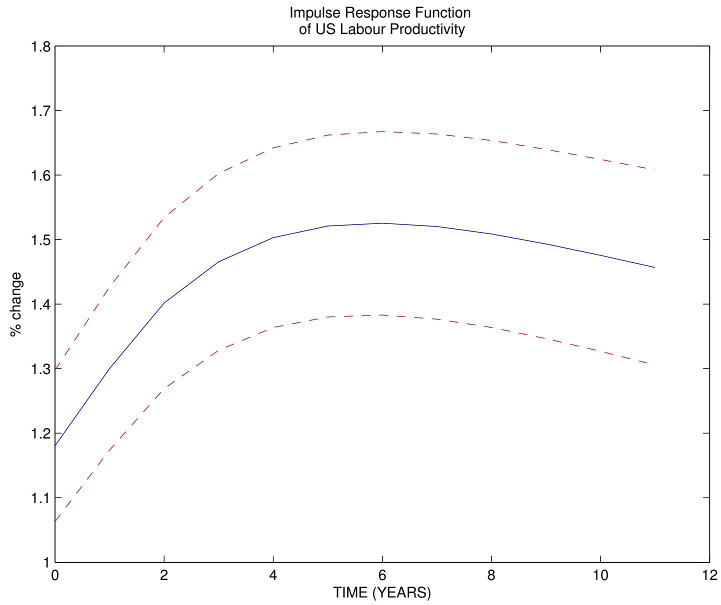

Figure 1: Accumulated IRFs of US Productivity Growth to the World Technology Shock. Dashed lines represent 95% confidence intervals.

US-EA productivity differential appreciates the dollar by approximately five percent[7]. Interestingly, IRFs here imply a similar result where a 0.8% decrease in the EA-US differential causes the euro to hit its lowest value (highest on the graph) at 4.0% lower than its equilibrium level. This observed Balassa-Samuelson effect is not expected in this setting since our identification method cannot distinguish productivity levels between sectors of tradeable and non-tradeble goods. However, Alquist and Chinn (2002) and Schnatz et al (2003) explain that productivity improvements can generate wealth effects which, if they induce a disproportionate increase in spending on local goods, can cause an appreciation of the real exchange rate. Clarida and Gali (1994) also find a US-dollar appreciation in response to a relative supply shock. What is more, the response is highly significant, with confidence intervals being rather narrow, and long-lasting, since it takes approximately five years for the effect to die out. This is instructive of the role of common innovations to technology in explaining relative price differences across countries, and re-enforces the results of Alquist and Chinn (2002) and Schantz et al (2003) as well as the theoretical insight of currency appreciation following a productivity shock due to wealth effects.

[7]Schnatz et al (2003) find this number to be lower at $1.5 - 2.0\%$.

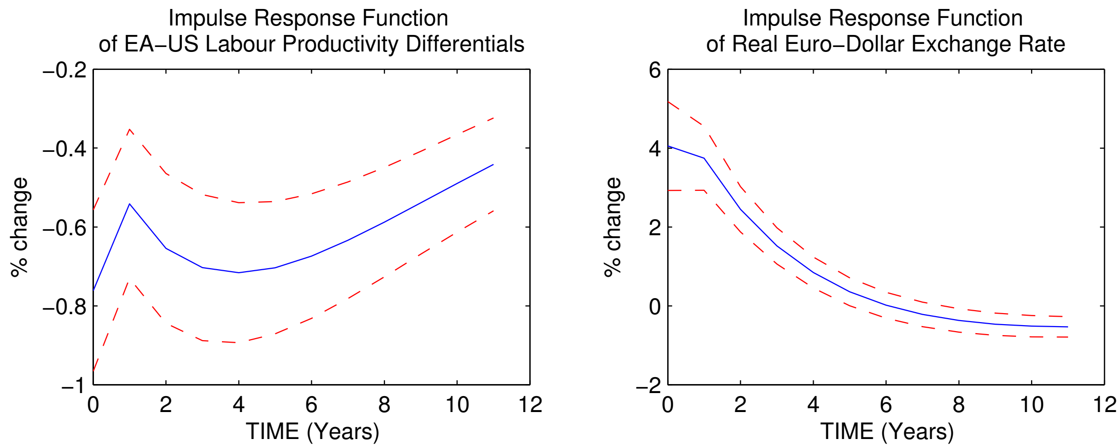

Figure 2: IRFs of log Productivity Differentials and the log of the Real Exchange Rate to the World Technology Shock. Dashed lines represent 95% confidence intervals.

We now move on to the analysis of the influence of the world technology shock to the variability of the underlying economic aggregates, as measured by its percentage contribution to the variance of the forecast error. The results are reported in Table 1. The biggest portion of the variability in the US productivity series in the short-run is explained by the permanent world productivity shock; 75.8% on impact and more than 90% at horizons of 5 years or above. Furthermore, the world technology shock explains one-fifth of the variation in productivity differences across the Atlantic at business-cycle frequencies and one-quarter at long horizons. The significance of the common technology innovation in explaining variations in productivity is in line with IRBC models that find significant contribution of the common factor to business cycle fluctuations (Canova et al 2007, Kose et al 2003, Stock and Watson 2005).

Moving on to the relative importance of the shocks to the real euro-dollar rate, we observe that permanent technology improvements explain 15.7% of the variance in the forecast-error on impact, and around 20% at business cycle frequencies. This shows that permanent innovations to world technology, identified by extracting a common stochastic trend in productivity, is an important source of fluctuations in relative prices; despite not being the dominant one. Ahmed et al (1993) show that a world supply shock explains 11.8% of the 32-quarters variation of the growth in relative prices. In their identification method the world supply shock is allowed to permanently change both the level of output and hours, thus making it similar to the disturbance identified in this study. Even though their findings are quantitatively lower than ours, care should be taken in their comparison. First, Ahmed et al (1993) use a six-variable SVAR system that allows for a much richer shock structure. Second, the effects concern the change in the real exchange rate, rather than the level, and last but not least their results concern a measure of relative price between US and a five-nation OECD aggregate, whereas we focus on the real euro-dollar rate. Moreover, the share of real exchange rate movements attributed to the world technology shock stands much higher than the supply shock identified in Clarida and Gali (1994) and the economy-wide productivity shock identified in Lewis (2006), both standing between $0-5\%$ at business cycle frequencies. The latter implies that indeed our identification method that allows for open-economy dynamics via cointegration, uncovers sources of movements largely omitted in the conventional identification of country-specific supply or productivity shocks. However, the findings of Farrant and Peersman (2006) mitigate the strength of this argument since, by repeating the exercise of Clarida and Gali (1994) using sign restrictions for identification, they find that shocks to the labour input explain between 12-28% of fluctuations in relative prices between EA and US, a number very close to our estimation. Finally, the contribution of the world technology shock is much lower when compared to shocks on (real) demand and nominal shocks (Lewis 2006, Clarida and Gali 1994, Farrant and Peersman 2006, Eichenbaum and Evans 1995). Our results do not stand in sharp contrast with these studies since the residual variability, which constitutes the biggest portion (around 80%), is explained by shocks which can be either real or nominal in nature, but cannot change productivity permanently. Lastrapes (1992) finds an important role of “real” shocks to real and nominal exchange rates; ranging between $60\% - 90\%$ at horizons of 1-60 months. However, Lastrapes (1992) approach cannot disentangle neither a supply, neither a technology nor a real demand shock. His method can only distinguish between a “real” and a “nominal” shock and Lastrapes himself acknowledges that if the world is subject to more than a single real shock his results can be potentially compromised. Indeed, this is why our estimates of the effect of permanent common innovations to technology are much lower than Lastrapes (1992).

Table 1: Variance Decomposition of the Endogenous Variables

Years	Contribution of the World Technology Shock to the variance of US Productivity (percent).
0	75.8 (27.6)
1	85.8 (27.8)
2	89.1 (26.8)
5	91.4 (22.0)
8	93.4 (16.5)
∞	100.0 (1.9)

Years	Contribution of the World Technology Shock to the Variance of EA-US Productivity Differentials (percent)
0	20.3 (22.9)
1	18.5 (22.0)
2	19.4 (22.5)
5	22.7 (23.6)
8	24.4 (24.0)
∞	25.9 (24.5)

Years	Contribution of the World Technology Shock to the Variance of the Real Euro-Dollar Exchange Rate (percent)
0	15.7 (17.8)
1	19.8 (18.2)
2	20.6 (18.2)
5	20.3 (17.9)
8	20.1 (17.8)
12	20.2 (17.9)
∞	20.4 (18.9)

Note: Numbers in brackets represent standard deviations.

Overall, our results suggest that if we are indeed identifying a “pure” technology shock, despite it being the primal source of productivity movements in US and very influential for productivity differentials in the two regions; it fails to be the primal source of real exchange rate variations. This provides limited support to productivity-based models of real exchange rate determination à la Balassa (1964) and Samuelson (1964). These type of models instruct that in a world with tradeable and non-tradeable goods, where the law of one price holds and under full capital mobility; only sectoral productivity differentials can drive, or even cause, movements in the real exchange rate (DeGregorio and Wolf 1994, Froot and Rogoff 1991, Obstfeld 1993, Rogoff 1992).

Despite the above, the effect of world productivity disturbances is in no case negligible and can account for 15%-20% of the euro-dollar real exchange rate movements at business cycle frequencies. Moreover, analysis of impulse responses has shown that common changes to productivity can cause significant and persistent deviations from PPP during the countries' transition to the new productivity levels; even if these changes are not concentrated on a particular sector as the Balassa-Samuelson framework instructs. The latter reinforces theoretical explanations of the direction of relative price movements in response to productivity shocks, as analysed and explained by Alquist and Chinn (2002), Schnatz et al (2003) and Bergstrand (1991). Therefore, identifying world technology shocks using the one-to-one cointegrating relation of country-productivities allows for the exploitation of open-economy dynamics and international spillovers that could serve as an important source of variation over the cycle. In that sense, our results meet with many different branches of the literature. As explained, they meet with the empirical SVAR literature on real exchange rate movements in so far that non-permanent technology shocks constitute the main driving force. Also, they build on the more theoretical work of Rabanal et al (2008) and Fève and Dupaigne (2009) who make use of the $(1-1)$ cointegrating relation in productivities in a general equilibrium framework. In particular, Rabanal et al (2008) show that this cointegrating relation implies higher persistence and slower productivity spillovers; the open-economy dynamics mentioned above which, as they explain, can imply more persistent and volatile exchange rates. As a result, their calibrated two-country DSGE model that uses a bivariate VEC structure of TFPs as the exogenous driving force can match very well the variability of the real exchange rate relative to output; compared to similar models with stationary, albeit persistent, TFP processes (Backus, Kehoe and Kydland 1994, Heathcote and Perri 2002). In a similar vein, Fève and Dupaigne (2009) stress the importance of a common stochastic trend in country-specific productivity series, interpreted as a "world technology shock", on uncovering open-economy dynamics important for explaining country-specific cyclical fluctuations. Last but not least, our results meet with a large number of dynamic factor models employed in multi-country settings that find a significant contribution of the common factor to business cycle fluctuations (Canova et al 2007, Kose et al 2003, Stock and Watson 2005).

4.1 Discussion

Interpretation of the results in this empirical work needs to be done with caution. The real exchange rate literature has hard time to assign an important role to supply shocks in explaining movements in relative prices, even though these type of disturbances account for 100% of the long-run variations in output. The estimated effect of world technology disturbances to real exchange rate movements, compared to their labour-supply counterparts is suggestive that real factors, whereby real I mean changes in underlying economic fundamentals that can permanently change the level of output, productivity or both, can potentially be a significant - albeit not the dominant - source of real exchange rate fluctuations. Nonetheless, non-permanent technology shocks - let them be temporary technology shocks, nominal shocks, demand shocks but as well as shocks to the labour input - appear to be the dominant determinants of relative prices.

Finally, one should not forget that the structural disturbance we attempt to extract in this study is a *common* world technology shock, where by "common" we mean a common stochastic trend in the country-specific productivity series that constitutes the only source of unit root to productivity. Intuitively, a common shock in a (two country) world where countries are symmetric and prices are fully-flexible should not cause any movements in the exchange rate. Consequently, an RBC-type of model would be expected to assign a very low role to a common world technology shock in moving relative prices, even when this disturbance is the only driving force of the cycle. For the latter, it is natural that the estimated effects of country-specific shocks are in general more relevant[8] than common disturbances are. However, what

[8]See Ahmed et al (1993). We also identify country-specific technology shocks by relaxing cointe-

we claim to identify here is an *additional* source of movements, largely omitted so far in the literature. A common technology disturbance can cause variations in the relative price of goods across two regions for one of two reasons. First, because countries are not symmetric and absorb technology differently[9] and second because there are frictions in the world economy causing prices not to adjust fully - at least not instantly. The latter effect though would be expected to dissipate in the long-run where prices are fully flexible. Therefore, the small reaction of the real exchange rate to the shock might be due to the fact that US and EA are able to absorb technological innovations in a similar manner and speed; rather than a weak relation between productivity and exchange rates. If that holds true, we should expect to observe a significant contribution of country-specific, temporary technological disturbances to the forecast error variance but this goes beyond the scope of this paper.

5 Robustness

In this section we evaluate the robustness of our results to different modifications. In particular, we check the sensitivity to changing the first variable of the system to the EA productivity growth, to adding another endogenous variable in the system, to the definition of price and hence the real exchange rate, to changing the frequency of the data and finally to relaxing the cointegration assumption. Overall, our results are robust in the sense that the findings for the contribution to variance are similar, if not higher, under different specifications with the exception being when we allow for permanent effects to the real exchange rate with quarterly data. From now on whenever the benchmark case is referred to, we mean the SVAR system analysed in the previous section.

gration as part of our robustness checks.

[9]Note that when assuming $(1-1)$ cointegrated productivities we are assuming that the countries are indeed symmetric at the steady-state.

5.1 EA Productivity Growth

In this first robustness test, we check whether our results change when we include EA productivity growth instead of US as the first variable. If we are indeed identifying a common productivity shock the results should not be altered dramatically. Also, Fernald (2007) shows that impulse responses in VAR systems with long-run restrictions can be quite sensitive to low-frequency correlations that need not be causal. Thus, to the extend that non-causal low-frequency correlations between the two productivity-growth series and the exchange rate are different, we might expect to get different results. If, however, this statistical feature with no economic value does not influence much the impulses (and hence the contributions to variance) the results should not be very different under the two specifications.

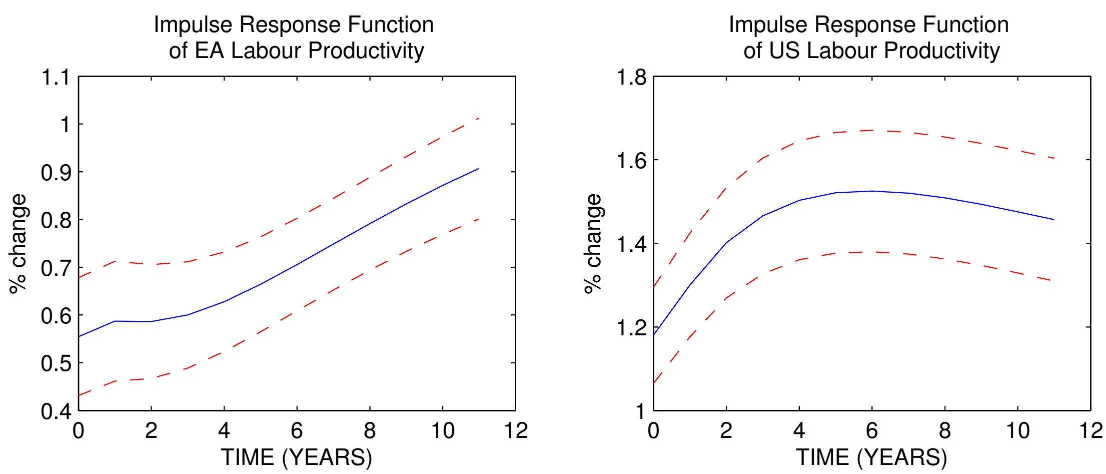

Figure 3: Accumulated IRFs of EA and US Productivity Growth to the World Technology Shock. Dashed lines represent 95% confidence intervals.

The response of EA productivity growth is instructive about the commonality of the shock; as representing a common stochastic trend in the productivity series. The impulse causes a positive and significant impact effect on EA productivity (figure 3) increasing it by 0.55%. Thereafter, the series increases monotonically until it is restored to its new long-run level approximately 1.2% higher than before the shock; matching exactly the permanent increase on US productivity (by construction).

Table 2 provides a summary of results on variance decompositions under different attempts. For this case - Panel B, point 1 of the table - we observe that the contribution of worldwide technological innovations at short horizons are much less influential in Europe than in the US (24.4% up to one year compared to 85% in the US at the same horizon) but are still quite significant. This implies that indeed US drives the world business cycle - a result found in Giannone et al (2008) - or that Europe is more affected by domestic disturbances.

Finally, the contributions to the real exchange rate are in general higher - see panel A of table 2 under specification 1. On impact, the world technology shock contributes 27.8% and 26.2% in twelve years. This is a bit different than in the benchmark case but shows that the potential effect of common permanent technology shocks to relative prices can be more than 25%. This discrepancy, albeit not being destructive raises some warning flags. In view of the latter, we thereafter check the stability of our results to a 4-variables SVAR system.

5.2 Four Variables

We include the "short-term interest rates differential" as an additional variable to our initial specification since this can possibly capture differences in monetary policy behaviour across the two regions. Clarida and Gali (1994), Farrant and Peersmann (2006) and Eichenbaum and Evans (1995) have shown that monetary policy shocks are very influential to real exchange rate movements. Further, interest rate differentials have a theoretical appeal in exchange rate determination (Frankel J. 1979).

We consider short-term interest rate differentials between US and Germany, instead of using some aggregate measure for Europe. The reason for doing so is twofold. First, common monetary policy for Europe does not exist before 1999, thus if we want to capture relative monetary policy shocks some aggregate measure of the interest rate would not make much sense. Second, Germany has always been a dominant country for fighting inflation in Europe. Clarida et al (1998) document that Bundesbank played a very influential role on monetary arrangements" within France, Italy and England. Also, it seems that these countries were following very closely German monetary policy even before the "hard ERM" period, as defined by the authors to be the

Table 2: Variance Decomposition Under Different Specifications

Panel A: Contribution of the World Technology Shock to the Variance of the Real Exchange Rate (percent)[1]				
Specification	0	1	5	12
1. EA Growth	27.8	28.6	26.9	26.2
	(18.9)	(18.6)	(18.1)	(18.1)
2. Four Variables	17.9	20.1	17.1	16.8
	(23.9)	(24.6)	(24.1)	(19.5)
3. Price of GDP RER	20.4	23.8	25.5	25.5
	(19.1)	(19.4)	(19.7)	(19.8)
4. Quarterly data - RER first-diff	0.9	0.3	0.6	1.9
	(3.3)	(3.3)	(4.2)	(6.6)
5. Quarterly data - RER in levels	8.6	10.8	13.7	11.9
	(14.3)	(15.2)	(15.5)	(14.6)
Panel B: Contribution of the World Technology Shock to other selected Variables				
Variable	0	1	5	12
1. EA productivity	23.5	24.4	46.0	90.3
	(30.5)	(31.3)	(26.8)	(12.2)
2. US productivity under 4 variables	65.8	81.4	85.1	94.3
	(23.8)	(27.8)	(23.8)	(11.2)
Panel C: Country Specific Permanent Shocks - No Cointegration				
Shock	0	1	5	12
US shock on RER	61.8	65.5	66.8	66.7
	(27.8)	(27.1)	(27.3)	(27.4)
EA shock on RER	45.9	45.5	46.4	46.6
	(27.7)	(26.7)	(26.5)	(26.6)

Note: Panel A: Specification one corresponds to the SVAR system where the first endogenous variable is Euro-Area productivity growth. Specification two corresponds to an SVAR system of 4 endogenous variables where the additional variable is interest-rate differentials, specification three to the estimation where the exchange rate was GDP-deflated and specifications four and five correspond to estimations using quarterly data. Under specification four, the real exchange rate was included in first-differences in the estimation and one lag was used whereas in case five the real exchange rate was included in levels and four lags were used. For specifications four and five the horizon is quarters, not years.
Panel B: The first line refers to the contribution of the world technology shock to EA productivity and the second to the contribution of the shock to US productivity in the four-variable SVAR.
Panel C: Effects of country-specific technology shocks on the real exchange rate when cointegration is relaxed. Numbers in brackets represent standard deviations.

years between 1990 – 92 (Clarida et al 1998).

Following Eichanbaum and Evans (1995) and Faust and Rogers (2003) we proxy the short term interest rates using the "Federal Reserve Overnight Effective Rate" for US and the "Money Market Frankfurt Banks Overnight Rate" for Germany; both at annual frequency[10]. The differentials series is stationary[11] and is included in levels in our system. Note also that we use Germany-US thus an increase in the series represents a contractionary monetary policy shock in the EA. We also check with a lower frequency measure of the short term interest rates across the Atlantic, namely the equivalent to the 3-Months rates to the above, again transformed at annual frequencies. Results (not reported) are not significantly different.

Figure 4 presents the response functions of the interest rate differential and the log-level real exchange rate to the world technology shock; for the sake of brevity we omit the dynamic responses of the other two variables. Overall, the responses maintain the same sign but we observe some minor differences in the magnitude when compared to the initial specification.

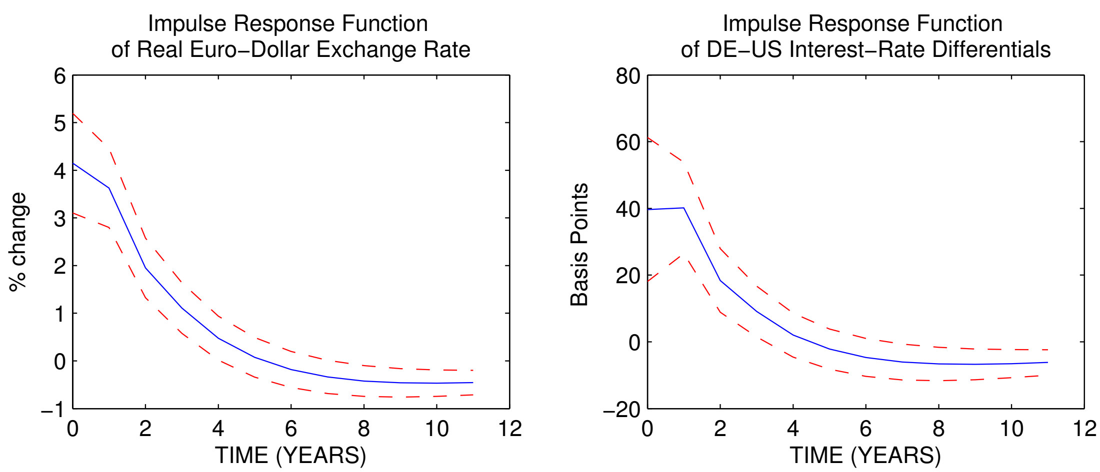

Figure 4: IRFs of Interest Rate Differentials and Euro-Dollar Real Exchange Rate to the World Technology Shock. Dashed lines represent 95% confidence intervals.

The effects of a world technology shock in explaining movements in real exchange rate are very similar compared to the benchmark case (see Table 2 - Panel A under specification 2). On impact,the contribution to real exchange rate movements is increased slightly, at 17.9% compared to 15.7% initially found and at business cycle frequencies the effect is close to 20% as in the initial specification. Hence, our conclusions are robust to the addition of interest rate differentials as an additional and

[10] Our indicator for Germany is a bit different than these authors

[11] ADF test rejects unit root at 5% level.

potentially significant variable for explaining movements in exchange rates.

5.3 Different Measure for the exchange rate

In this section, we retain the exact same specifications for the first two endogenous variables of our system as in the benchmark case, but we use a different measure of the real exchange rate. Using different price indicators to construct the exchange rate can be important especially if there are significant differences in their composition of traded and non-traded goods (Bets and Kehoe 2006). With that in mind, we use the "Price of GDP" from the same sources - the AWM of the ECB expanded for the missing years for the EA and data from the BLS for the US. Figure 5 which plots both exchange rate series depicts some discrepancies at the beginning and the end of our sample. Stationarity results are still quite solid as in the benchmark case, with a p-value for the ADF test being 0.02.

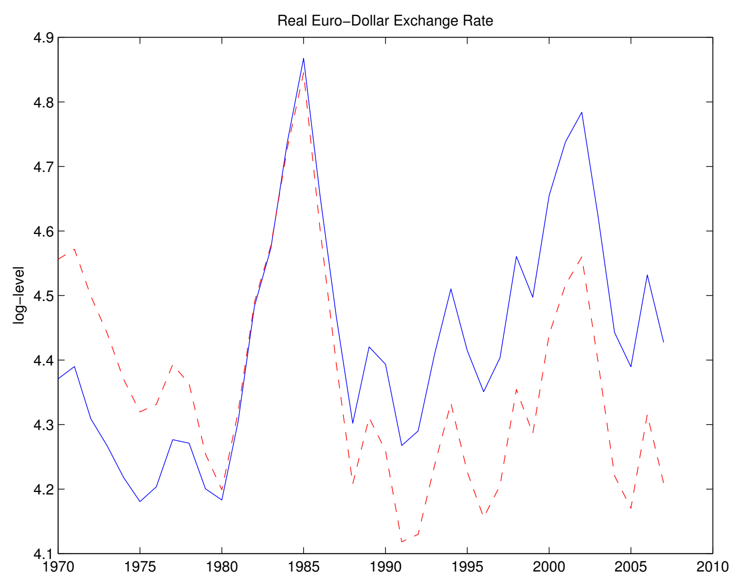

Figure 5: Log real exchange rate, GDP deflated (dashed line) and HCIP deflated (solid line)

IRFs are very similar so they are not analysed in detail for the sake of brevity. On variance decompositions, the contribution of the world technology shock on the real euro-dollar rate is elevated by almost 5% at all horizons (see Table 2, Panel A under specification 3). Thus, this test retains robustness while allowing for the possibility that permanent technology disturbances are a bit more important than initially estimated.

5.4 Quarterly Data

A natural criticism for our study is that we do not use data at quarterly frequency, something that can be at least partly justified by better quality of annual data. Nevertheless, in this section we check for the robustness of our results to the use of higher frequency data, at the cost of lower quality. In particular, this involves using productivity *per person* instead of the preferred measure of productivity per hour. Our sample period is 1970Q1 - 2000Q4. Over this period, and with quarterly data, the euro-dollar real exchange rate is found to be non-stationary[12] and at a first attempt it is included in first-differences in our SVAR. Many things are worth mentioning out of this analysis. First, we observe that the effect of the world technology shock on the real euro-dollar rate is negative and significant, albeit short-lived, implying a dollar depreciation (see figure 6). This stands in contrast to the corresponding response in the benchmark case, but as well as to evidence of a Balassa-Samuelson effect on the euro-dollar rate documented in the literature (Alquist and Chinn 2002, Schnatz et al 2003). Second, our results are not immune to these kind of changes, with the contribution of the world technology shock to the variability of the log-level real euro-dollar rate standing between 0.9%-3.4% (Table 2, Panel A, Specification 4) at horizons between one to twenty quarters (five years); much lower than the benchmark case.However, there is a problem of identification under this specification: the world technology shock identified with quarterly data and the real exchange rate in first differences is effectively a US-specific technology shock. That is, if we extract such an innovation by relaxing cointegration, the correlation between the two is almost perfect standing at 0.997(!)[13] .

Nevertheless, the discrepancy decreases a great deal once we abstract from a

[12] ADF tests give a high p-value, 41%. There are evidence that real exchange rates have been more volatile and exhibit a unit-root in the post Bretton Woods period, see for example Clarida and Gali (1994). A graph of the real exchange rate series at quarterly frequency is given in Appendix A.

[13] We analyse the case where we relax cointegration in more detail in the following section.

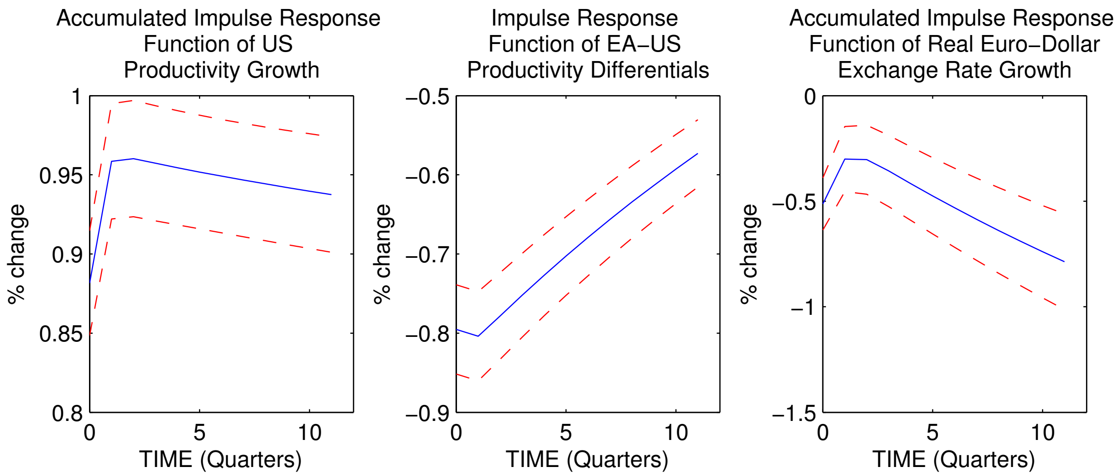

Figure 6: IRFs of the endogenous variables to the World Technology Shock, quarterly data. Dashed lines represent 95% confidence intervals.

unit-root in the real exchange rate even at quarterly frequency[14]. Specifically, the contribution of the shock to real exchange rate variance is higher ranging between 8.6% on impact, 11.4% at twenty quarters (five years) and 15.3% in the long-run[15] (Table 2, Panel A, Specification 5). Compared to the benchmark case the contribution is still lower but not vastly different. Importantly, the correlation between this shock and a US-specific one is much lower than before, with a correlation coefficient equal to 0.70, and significantly different from one[16].

Overall, we can derive one conclusion out of this exercise: our results are not very sensitive to the use of quarterly or annual data to extract the world technology shock, but we cannot say the same thing about how the real exchange rate is specified. If we assume the real exchange rate as an I(1) process, we find that the effect of common shocks to technology are not important for relative prices across the Atlantic. However, if the real exchange rate is I(0), their contribution is in no case negligible but a bit lower than in our benchmark specification.

[14] Note that even though this might not be common practice, CADF tests do provide some evidence of stationarity over our sample period.

[15] We estimated the VAR with 4 lags in order to "absorb" some of the persistence in the level of the real exchange rate.

[16] We bootstrapped 1000 random samples of correlations, to derive standard errors in order to test the null hypothesis that the correlation between the two shocks is perfect. The latter is easily rejected, with a t-value equal to -178.74.

5.5 Relaxing Cointegration

In this section, we relax the assumption of cointegrated productivities, thus allowing for the possibility of permanent, country-specific technological innovations and consequent productivity improvements. The SVAR system is estimated by using productivity growth of US and EA (log-first-differences) and the real exchange rate in levels:

$$\begin{pmatrix} \Delta X_t^{US} \\ \Delta X_t^{EA} \\ q_t \end{pmatrix} = C(L) \times \begin{pmatrix} \epsilon_t^{US} \\ \epsilon_t^2 \\ \epsilon_t^3 \end{pmatrix}$$

with:

$$C(1) = \begin{pmatrix} c_{11} & 0 & 0 \\ c_{21} & c_{22} & 0 \\ c_{31} & c_{32} & c_{33} \end{pmatrix}$$

X_t is labour productivity, Δ is the difference operator and q is the real euro-dollar rate, all expressed in logs. The first structural disturbance can be identified as a US-specific technology shock, since it is the only shock that can permanently change the level of US labour productivity, whereas ϵ_t^3 can be identified as a "non-technology shock", since it constitutes the only disturbance in this empirical economy that cannot permanently change the level of either US or EA productivity. However, identification of the second disturbance is not clear. We could label ϵ_t^2 as a EA-specific technology shock that does not spill-over to the US, consistent with evidence that spillovers originate in the US (Giannnone et al 2008), but its effects can be blurred by the influence of ϵ_t^{US}. For this reason, we do not label this shock. The IRFs from this exercise are presented in figure 7. As found in many parts in the literature, we also obtain a significant dollar-appreciation following a US-specific shock.

The results of this exercise are pretty interesting. First, we observe that the correlation of the two technology shocks, the US-specific shock extracted from this system and the world technology shock, is positive with a correlation coefficient equal to 0.795[17]. This supports the idea that innovations to world technology, represented as a common stochastic trend in country-specific productivity, include a prominent US component. Moreover, it supports the purpose of this paper as identifying an additional source of movements so far largely omitted in the literature; since the correlation is far from being perfect[18]. Secondly, the results point that US-specific technology shocks account for the bulk of the movements in the real euro-dollar rate, contributing around two-thirds to its variability at business-cycle frequencies (Table 2, Panel C, "US Shock on RER"). A EA-specific shock identified in a similar manner, that is estimating the system as above but with the EA being the first variable in the Choleski ordering and US the second, has lower but still important contribution to exchange rate variability ranging around 45-47% at all horizons[19]. This is interesting since technology shocks seem to be very important for explaining real exchange rate movements, supporting the RBC paradigm and productivity-based models of the real exchange rate. Also, US-specific technology shocks alone are the dominant driving force of relative price movements, counter to the studies of Clarida and Gali (1994), Farrant and Peersman (2006) and others who support that demand shocks or nominal shocks play a dominant role. The reader should notice that under our identification method, the third shock "nests" all those shocks that do not change productivity, including permanent shocks to the labour input, demand or nominal shocks and shocks to government purchases. In our estimation this shock contributes to around one-fifth of movements in relative prices at all horizons. Finally, own-country technology shocks have a higher effect than the world technology shock, a result that is not surprising based on the fact that the commonality of innovations to world technology causes more symmetric responses of prices in the two countries and

[17]Correlation is 0.234 if we identify a EA-specific technology shock in a similar manner.

[18]We bootstrapped 1000 random samples of correlations, to derive standard errors in order to test the null hypothesis that the correlation between the two shocks is perfect. The latter is easily rejected, with a t-value equal to -95.56.

[19]See Table 2, Panel C, "EA shock on RER". Note also that a second endogenous variable is not necessary for identifying a country-specific technology shock, that is we could have obtained the same innovation from a bivariate system $(\Delta X_t^{US} q)'$ and impose Blanchard and Quah (1989). Applying the latter leaves the effect of permanent, country-specific technological innovations to the real exchange rate unchanged.

consequently a less variable response of the real exchange rate.

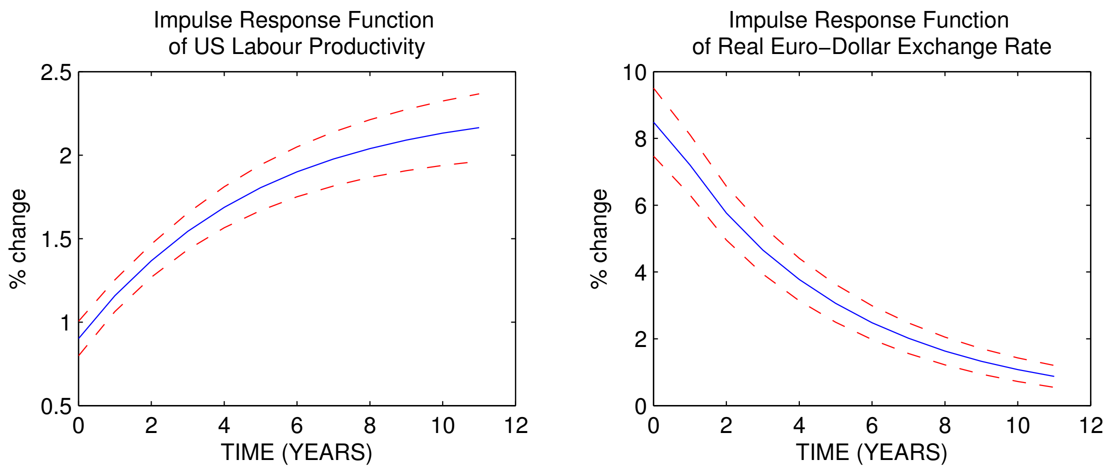

Figure 7: IRFs of the endogenous variables to a US-specific Technology Shock. Dashed lines represent 95% confidence intervals.

However, there are some issues in the estimation of the previous model worth discussing, and the main issue comes from the fact that when using quarterly data, and more specifically including the growth in the real exchange rate in our SVAR; improvements to productivity do not seem to play such an important role. Again, using the sample 1970Q1-2000Q4 the effect of US-specific technological innovations on the real exchange rate ranges only from $0.8\% - 1.3\%$ at all horizons. Yet again, the long-run matrix estimated out of quarterly data again points to "non-technology shocks" as the sources of a unit root in the real euro-dollar rate, providing more support to the results of Clarida and Gali (1994) and their followers, as well as to models that emphasise the role of non-technology shocks, let them be real (shocks to government purchases, labour input, real demand) or nominal (monetary shocks). This is not counter-intuitive since, as emphasised earlier, the definition of "non-technology" shocks nests many types of structural disturbances many of which are real in nature and can cause permanent changes to output. The latter is recognised by Clarida and Gali (1994) themselves who acknowledge that "real shocks to supply and demand account for more than 50% of the variance in forecasting real exchange rates".

However, the aforementioned results can be totally reversed once we assume that the real exchange rate is a stationary process, as instructed by the theory of Purchasing Power Parity that is employed by many models of the business cycle in open economies. This involves estimating a SVAR using $(\Delta X_t^{US} \Delta X_t^{EA} q_t)'$ with quarterly data to extract the shocks. In that case, the US-specific technology shocks account for the bulk of the movements in the real euro-dollar rate, contributing 46.8% of the variability in the same-quarter forecast error and between 60%-67% at horizons between 5-32 quarters[20]; numbers that are not so different from the annual frequency results.

Overall, this section shows that if the world is characterised by technology improvements that are country-specific, their contribution on the variance of the forecast error in the real exchange rate is potentially higher than a common stochastic trend in productivity. Moreover, it can possibly constitute the dominant source of fluctuations in relative prices at business cycle frequencies. However, as was the case in the previous sub-section, the importance of country technological innovations in explaining movements in real exchange rates is not sensitive to the use of quarterly data as such but rather to the specification of the real exchange rate series.

6 Conclusions

This paper adds to the empirical SVAR literature which looks for the underlying economic disturbances that are the prime sources behind real exchange rate fluctuations, sparked primarily by the seminal paper of Clarida and Gali (1994). Prominent examples include Farrant and Peersman (2006), Ahmed et al (1993), Lewis (2006); as well as some earlier work by Eichenbaum and Evans (1995) and Lastrapes (1992). We add to this list a consideration of an additional source of movements, namely a common stochastic trend in productivity interpreted as a world technology shock. In other words, we assume that productivities are one-to-one cointegrated; an assumption that is verified in the data using statistical methods. Our work is inspired by Rabanal et al (2008) and Dupaigne and Fève (2009) who have stressed the importance of this common trend in uncovering open-economy dynamics significant for

[20]Using 4-lags in the reduced-form VAR to absorb some of the exchange rate persistence.

cyclical movements in relative prices; as well as other macroeconomic aggregates. We also claim that this method of identifying permanent changes in technology is more immune to stationary country-specific shocks.

We find that the dollar appreciates significantly in real terms in response to the shock, suggestive of the importance of productivity differentials in explaining movements in relative prices. The contribution of the world technology shock to the variance of the forecast error in the real euro-dollar rate stands around 20% at business cycle frequencies. This does not contrast the existing literature that finds that demand or nominal shocks constitute the prime source of real exchange rate movements and provides limited support to productivity-based models of real exchange rate determination. However, even though common innovations to technology fail to be a dominant source of swings in real exchange rates, their effect is in no case negligible. Future work should shed more light on the relation between technology shocks and real exchange rates, as well as the proper identification of these innovations in SVARs. Obviously, clearer evidence on the latter will help in answering the former.

6.1 Acknowledgments

I thank Patrick Fève, Franck Portier, Martial Dupaigne, seminar participants in Toulouse, at the Doctoral Conferences in Bochum and Montpellier (2010) and the ASSET 2010 meeting at the University of Alicante for useful comments and discussions.

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7 Appendix A

Real Exchange Rate: The series of the euro-dollar real exchange rate was constructed using values of the nominal "synthetic" euro, provided by Datastream and consumer prices indices from the AWM for the Euro Area and the Bureu of Labour statistics or the US. The AWM Harmonised Consumer Price Index (HCIP) provides aggregate price data for eleven European countries at quarterly frequency, seasonally adjusted. The series was converted to annual frequency by simple averaging of the quarterly data. The US price series used concerns US-city average HCIP data, non-seasonally adjusted. The base of the real exchange rate index are the years 1982-84. GDP deflated real exchange rate used price series from the same sources.

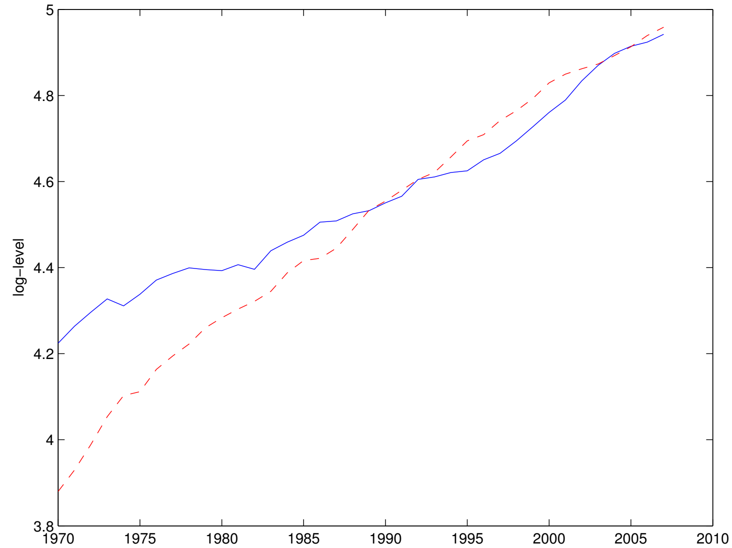

Figure 8: Log-US Productivity (solid line) and log-EA Productivity (dashed line).

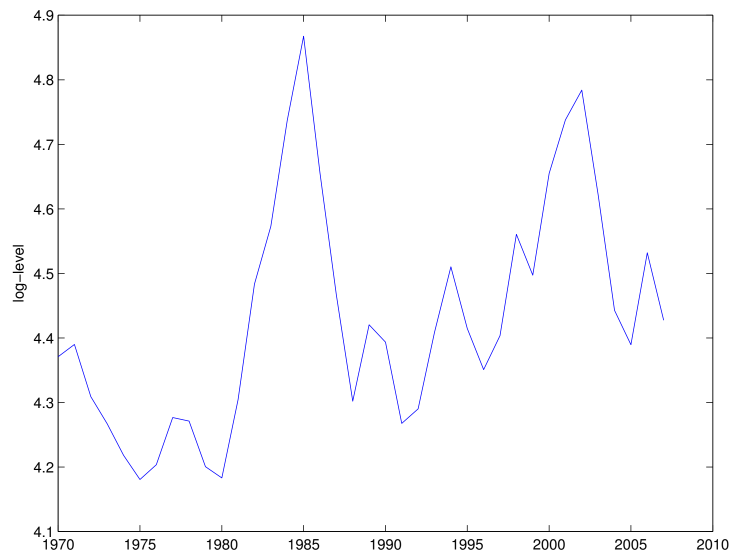

Figure 9: Data - Log Real Exchange Rate

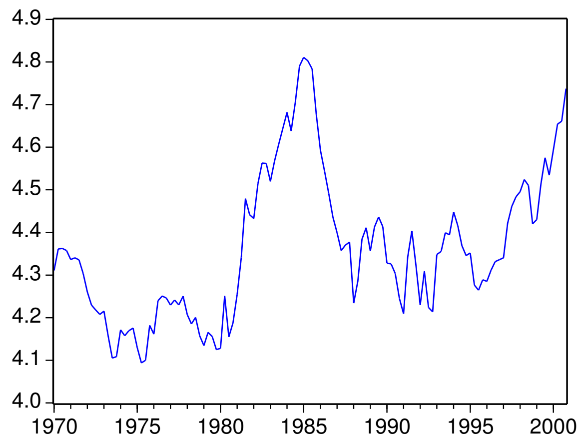

Figure 10: Data - Log Real Exchange Rate, quarterly frequency

8 Appendix B

Table 3: Stationarity Tests for the log EA Productivity

	Test		
	ADF[1]	Phillips Perron[2]	KPSS[3]
Null	Series	Series	Series
Hypothesis:	has a unit root	has a unit root	is Stationary
t-Statistic	-2.560381	0.469001	0.749793
Test Critical value			
1%	-3.626784	-3.621023	0.739000
5%	-2.945842	-2.943427	0.463000
10%	-2.611531	-2.610263	0.347000
p-value	0.1104	0.0073	-

Tests include an intercept. [1]One lag in first-differences.
[2,3]PP and KPSS bandwidth were set to 2 and 5 respectively, as per the Newey-West bandwidth.

Table 4: Stationarity Tests for the log US Productivity

	Test		
	ADF[1]	Phillips Perron[2]	KPSS[3]
Null	Series	Series	Series
Hypothesis:	has a unit root	has a unit root	is Stationary
t-Statistic	1.142251	0.837019	0.738955
Test Critical value			
1%	-3.626784	-3.621023	0.739000
5%	-2.945842	-2.943427	0.463000
10%	-2.611531	-2.610263	0.347000
p-value	0.9971	0.9934	-

Tests include an intercept. [1]One lag in first-differences.
[2,3]PP and KPSS bandwidth were set to 0 and 5 respectively, as per the Newey-West bandwidth.

Table 5: Stationarity Tests for the EA-US log Productivity Differentials

	Test		
	ADF[1]	Phillips Perron[2]	KPSS[3]
Null Hypothesis:	Series has a unit root	Series has a unit root	Series is Stationary
t-Statistic	-3.235975	-3.300347	0.655865
Test Critical value			
1%	-3.626784	-3.621023	0.739000
5%	-2.945842	-2.943427	0.463000
10%	-2.611531	-2.610263	0.347000
p-value	0.0259	0.0221	-

Tests include an intercept. [1]One lag in first-differences.
[2,3]PP and KPSS bandwidth were set to 3 and 5 respectively, as per the Newey-West bandwidth.

Table 6: Stationarity Tests for the log Euro-Dollar Real Exchange Rate

	Test		
	ADF[1]	Phillips Perron[2]	KPSS[3]
Null Hypothesis:	Series has a unit root	Series has a unit root	Series is Stationary
t-Statistic	-2.797551	-2.200538	0.330870
Test Critical value			
1%	-3.626784	-3.621023	0.739000
5%	-2.945842	-2.943427	0.463000
10%	-2.611531	-2.610263	0.347000
p-value	0.0686	0.2095	-
	CADF Tests[4]		
Covariate	Lags	Leads	P-value
Productivity differentials	1	0	0.0148
Productivity differentials	1	1	0.0066
US Productivity Growth	1	0	0.0487
US Productivity Growth	1	1	0.0133

Tests include an intercept. [1]One lag in first-differences.
[2,3]PP and KPSS bandwidth were set to 3 and 4 respectively, as per the Newey-West bandwidth.
[4]CADF Tests include one lag of the real exchange rate in first differences.

Table 7: Cointegration Tests for EA-US Productivities

Tests' P-Values		
Number of cointegrating vectors	Trace	Max Eigenvalue
0	0.0103	0.0332
1	0.0832	0.0832
Restriction: $\beta = (1 - 1)'$		
Number of cointegrating vectors		P- Value
1		0.510002

Tests include a constant term in the cointegrating vector but no deterministic trend in the data. We used one lag in first differences.