Deviations in real exchange rate levels in OECD countries and their structural determinants

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Abstract

Real exchange rate, a key international relative price, has puzzled generations of economists. We show that, when productivity and labor wedges are precisely measured, an augmented Balassa-Samuelson theory can explain real exchange rate behavior of the majority of OECD countries accounting for nearly two thirds of the world GDP between 1970 and 2017. We carefully construct a new dataset of total factor productivity (TFP) by sector in levels to facilitate a meaningful cross-sectional inference. Levels of real exchange rate can be explained by differences in traded and nontraded TFP levels in the direction predicted by the Balassa-Samuelson hypothesis. However, this relationship is only significant after we account for the presence of differences in labor market structure that drive supply-side labor wedges. We discuss mechanisms for labor market structural differences to differentially affect real exchange rates in a currency union and in freely-floating exchange rate economies.

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1 Introduction

Real exchange rate is a key relative price in international economics. Yet, the empirical behavior of real exchange rates puzzles generations of economists. Recently, Berka et al. (2018) successfully explain both cross-sectional and time-series variations in real exchange rates in the eurozone, while highlighting the importance of labor wedges as a key additional driver to the Balassa-Samuelson-type models that are based on sectoral productivities alone (Karabarbounis (2014), Balassa (1964), Samuelson (1964)). Engel and Zhu (2018) also discuss the differences in exchange rate puzzles across countries with fixed and floating exchange rates and highlight the variation in statistical relationships beyond exchange rate "puzzles" across different exchange rate regimes.

Our contributions are both to the literature on real exchange rate determination and the literature on currency misalignment measurement. In an important contribution, Engel (2011) shows that when currencies are misaligned, optimal monetary policy targets not only a linear combination of output and inflation gaps (as found by Clarida et al. (2002) earlier), but also the exchange rate misalignment. A key factor limiting the exchange rate misalignment literature is the availability of data on price level differences. We partly synthesises these strands of literature by carefully constructing a comprehensive dataset that allows us to discuss real exchange rate level differences and their theoretically-relevant drivers across countries with different exchange rate regimes.

Specifically, we construct a dataset of total factor productivity ("TFP" thereafter) by sector, real exchange rates, and labor market institutional characteristics for 18 OECD countries that account for nearly two-thirds of the world GDP, starting in 1970s. We construct all variables in levels, so as to draw meaningful inference from both time-series and cross-sectional variation in the data. Our empirical results indicate that the cross-sectional variation, as well as the more precise measurement of productivity through TFP rather than labor productivity, are fundamental in generating empirical support for the Balassa-Samuelson hypothesis. We also show that supply-side labor wedges are key in further elucidating the Balassa-Samuelson explanation of real exchange rates, beyond productivity measures. We particularly highlight the role of unionization concentration variable as a driver of labor-market wedges, and its role be particularly strong in the eurozone.

Although a meaningful cross-sectional information about price differences is crucial in measuring and explaining deviations of real exchange rates from the purchasing power parity, most studies that investigate the link between real exchange rates and productivity solely focus on the time-series variation[1]. Such studies also typically use labor productivity, in spite of its well-known limitations.[2] However, modern understanding of the theoretical framework underpinning the Balassa-Samuelson relationship is based on the more exogenous TFP. In such models (Benigno and Thoenissen (2003), Fitzgerald (2003) and others that followed), an increase in the traded TFP is predicted to appreciate real exchange rate by raising real wages thus forcing firms in the nontraded sector to raise the relative price of nontraded to traded goods. Conversely, an increase in relative nontraded TFP depreciates real exchange rate in these models. The strength of this effect is guided by the assumption perfectly competitive labor markets, as well as by the elasticity of substitution between goods in the traded sector. However, the evidence mostly rejects Balassa-Samuelson hypothesis in time series domain, except in cointegration studies. While cross-sectional studies generally find stronger support for Balassa-Sameulson hypothesis, particularly when comparing rich and poor economies, these studies are not based on TFP but labor productivity.[3] An exception is Berka et al. (2018), who construct measures of sectoral levels of TFP and real exchange rates, and find support for a Balassa-Samuelson relationship for nine eurozone economies between 1995 and 2009, after controlling for differences in labor wedges. We allow for the possibility of sectoral labor wedges on the supply side as in Galí et al. (2007), Karabarbounis (2014), and Berka et al. (2018). We argue these require an additional empirical measure of sectoral labor market wedges. We construct a direct and an indirect measure of labor supply wedges to show these significantly improve the fit of the augmented Balassa-Samuleson model.

As far as we are aware, ours is the first paper to find robust evidence in support of an augmented Balassa-Samuelson model both in cross-section and in time-series, among floating exchange rate developed countries. In our data, union concentration measures play a key role in elucidating the Balassa-Samuelson relationship. Their role is particularly important in countries with a common currency. We show that our results survive a large battery of

[1] See, e.g., Canzoneri et al. (1996), Chinn and Johnston (1997), Choudri and Schembri (2014), De Gregorio et al. (1994), Lee and Tang (2007), Lothian and Taylor (2008), Tica and Družić (2006), Chong et al. (2012), Ricci et al. (2013), Gubler and Sax (2019), or Chong et al. (2012).

[2] Labor productivity confounds the effects of the total factor productivity with the intensities of capital-to-labor ratios, intermediate input intensities, skill intensities, and differences in industrial structure.

[3] For example, see Rogoff (1996), Bergin et al. (2006), Chen et al. (2015), and Berka et al. (2012).

robustness checks. By using all available sources of data to construct vintages of price and productivity measures, we show how vintages influence the results of our baseline regressions[4]. This can explain the sometimes contradictory findings in the literature on Balassa-Samuelson hypothesis.

Finally, we contribute to the literature on real exchange rate misalignment. We borrow from that literature and construct the average deviations in real exchange rates that cannot be accounted for the fundamentals in our model, and highlight a number of countries where large average unexplained RER deviations remain.

The rest of the paper is organized as follows. The next section describes the construction of our datasets. Section 3 outlines the predictions of a basic model. Section 4 outlines the empirical methodology and section 5 the results and various robustness checks. Section 6 concludes.

2 Description of the data

We construct a panel dataset of sectoral TFP levels, real exchange rate levels, levels of relative unit labor costs, terms of trade, and measures of structural labor market differences for 17 OECD countries. We choose Germany as our base country because it allows for a clearer discussion of the roles played by an exchange rate regime, as well as geographic proximity. Our sample includes 8 countries with floating exchange rates and 9 countries that adopted the euro in 1999 (most of whom had previously been part of a crawling-peg exchange rate arrangement), and one country with a pegged exchange rate[5]. Our unbalanced annual panel covers a period from 1970 to 2017[6]. Over this period of time, countries in our dataset account for approximately 66% of world GDP (excluding the US, this share is 38%).

Data Appendix **??** online provides full details of data construction, which mimics the approach in Berka et al. (2018). Using concordances, we construct a panel of annual estimates of TFP and real exchange rates by combining cross-sectional TFP and PPP levels for benchmark years to indices

[4]We use different weighting schemes, different coefficient assumptions, and alternative relative price measures.

[5]The countries are: Australia, Austria, Belgium, Czech Republic, Denmark, Spain, Finland, France, Germany, Hungary, Ireland, Italy, Japan, the Netherlands, New Zealand, Sweden, and the United Kingdom. Eight of these countries were amongst the founding members of the eurozone in 1999.

[6]The length of data varies from a minimum of 13 years to a maximum of 48 years. Tables **??** and **??** in the Appendix describes the data sources, starting and ending dates.

of industry productivity and prices. Specifically, industry TFP levels are constructed using Groningen Growth and Development Centre (GGDC) Productivity Level database (1997 benchmark year), and are expressed relative to Germany[7]. A logarithm TFP level vis-à-vis the US is then: $a_{i,j,t} = \log\left(\frac{TFPlevel^{97}_{i,j} \times TFPindex_{i,j,t}}{TFPindex_{i,GER,t}}\right)$ where $TFPlevel^{97}$ is the level of country i's TFP in sector j relative to Germany in 1997, and $TFPindex$ is the time-series index of TFP in sector j, normalized to 1 in 1997. We aggregate $a_{i,j,t}$ across 11 industries into traded and nontraded aggregates ($a_{i,T,t}$ and $a_{i,N,t}$ respectively) using constant 1997 gross value added (GVA) country-specific weights and a standard industrial classification.

Figure 1 plots the levels of traded and nontraded TFP for each country compared to Germany. In the unbalanced panel, the level of TFP in traded sector is the highest in the Belgium, Ireland, and the US, and the lowest in east Europe. TFP in nontraded sector is also the highest in the Netherlands and the US, while it is the lowest Japan, the Czech Republic, and Hungary. Figures 4 to 6 compare our estimates of TFP levels to labor productivity level estimates from Mano and Castillo (2015). In theory, TFP influences labor productivity directly, and also indirectly through its effect on the capital-labor ratio, and possibly on sectoral skill accumulation. The strength of the indirect effects depends on the form of a production function. Thus, while there are several possible explanations for the differences between TFP and labor productivity, we nevertheless find it instructive to plot them jointly for comparison. For many countries, the relative levels and trends correspond closely with those in relative labor productivity. But there are exceptions: in Austria, Japan, Hungary, and less so in Denmark, TFP is lower than labor productivity in traded sector. On the contrary, the Netherlands and Ireland see higher TFP than labor productivity levels in the traded sector in parts of the sample period. Traded TFP generally shows larger volatility than labor productivity. While many countries only see minor changes in their nontraded labor productivity when compared with Germany, we observe a decline in nontraded TFP in Belgium, Japan, Spain and Italy throughout the sample, and in the UK and the US in the 1970s. The ratio of traded to nontraded TFP relative to Germany is most notably higher than the labor productivity ratio in Japan, Belgium, Ireland, and France, and lower in Denmark, the Netherlands, New Zealand and the US. The highest growth rates of the relative TFP (traded to nontraded sectors) are in Italy, the US,

[7] For New Zealand, which is not included in the GGDC database, we instead use Mason (2013)'s 2009 year benchmark sectoral TFP comparisons between New Zealand and Australia. With Australia in the GGDC database, this allows us to express New Zealand figures relative to the US, and consequently relative to Germany.

and Belgium, while the lowest are observed in Finland, the UK, and New Zealand[8]. The correlation between labor productivity and TFP is positive: over 0.8 for the traded and around 0.4 for the nontraded sector.

Our panel of real exchange rate levels is constructed using bilateral nominal exchange rates and relative price levels. The logarithm of the level of bilateral real exchange rate of country i relative to the US is defined as $q_{i,t} \equiv NER_{i:GER,t} + p_{i,t} - p_{GER,t}$, where NER is the logarithm of the German price of one unit of domestic currency, so that an increase represents an appreciation. $p_{i,t}$ and $p_{GER,t}$ denote logs of aggregate consumer price levels in country i and Germany, respectively, and are obtained from the International Comparison Program (ICP) aggregate consumer price PPPs (2011 vintage). We construct traded and nontraded price levels using the ICP price parities and goods and services CPI series as proxies for traded and nontraded price time series.[9]

Tables 1, 2, and 3 report some stylized facts of our variables of interest. For most countries, gaps in traded TFP vis-à-vis Germany tend to be below those for nontraded TFP, as well as being negative on average. Traded TFP also tends to be more volatile than nontraded TFP. For RER, east European countries in our sample have the lowest level of the real exchange rate, while Denmark, Sweden and Finland the highest. The east European countries have seen the largest appreciation of their q, while Sweden and Belgium depreciated the most relative to the US. Hungary and Japan see the highest q volatilty, and the United Kingdom (UK) the lowest.

[8]Bertinelli et al. (2016) produce labor productivity growth rates for traded and nontraded goods for a selected group of OECD economies using EU KLEMS for a balanced panel of 1970-2007, relative to the US. Their estimates suggest that relative labor productivity grew the fastest in Ireland, Finland and Spain, and slowest in Germany, Australia and Denmark.

[9]Frequently, studies use value-added deflators when constructing the price indexes of traded and nontraded goods (e.g., Drozd and Nosal 2010, Mihaljek and Klau 2008, Mihaljek and Klau 2004, Engel 1999). Real exchange rate is often measured as an index without a meaningful cross-sectional dimension (e.g., Bordo et al. 2017, Chong et al. 2012, Gubler and Sax 2019, Ricci et al. 2013). Papers that use value-added-based relative price measures tend to find a positive relationship between relative sectoral prices and real exchange rates (see Steenkamp 2013 or Drozd and Nosal 2010). However, the use of such value-added-based price indexes may bias results towards the acceptance of the Balassa-Samuelson hypothesis by introducing spurious correlation into the Balanssa-Samuelson regression. In our sample, the time-series correlation between sectoral TFP measures and value added-based price indices is higher than for consumer price based indices (see Figure **??** in the Appendix). Similarly, we observe that producer-price based indexes exhibit different sectoral inflation rates on average, especially for traded prices (see Steenkamp 2013 and Figure **??** in the Appendix).

In addition to the productivity variables, Berka et al. (2018) show that relative unit labor cost measures proxy for supply-side labor wedges in an augmented Balassa-Samuelson type model. We construct relative Unit Labor Cost levels (ULC) from OECD data, expressed as the average unit labor cost in country i relative to the unit labor cost in Germany after converting them into the same currency. To remove the mechanical influence of nominal exchange rates on relative ULC, we also construct relative Orthogonalized Unit Labor Costs ($OULC$ thereafter) that are orthogonal to the NER variation[10]. Table 1 and Figure 2 show that the lowest relative unit labor costs on average were found in the Spain, Italy, and Ireland, especially in the 1970s and 80s. More recently, the lowest unit labour costs are observed in Czech Republic and Hungary. The highest are in Japan and Austria.

We measure terms of trade ($TOT_{i,t}$) as the difference between export and import price levels from Feenstra et al. (2015), who construct them as export and import PPPs divided by the nominal exchange rate, relative to the base country. As with the other variables, they are expressed in logs. Finally, we construct bilateral long-run real interest rate differentials relative to the Germany ($RIRDIFF_{i,t}$) using the 10-year government bond yields obtained from Bloomberg. Relative interest rate levels are expressed as the home country rate less the German rate, adjusted by relative CPI inflation rates. Over the full sample, real interest rates are the highest in Denmark, and the lowest in Japan.

2.1 Institutional labor market differences

As has been appreciated since at least Leontief (1946), measures of union density or the ability of unions to coordinate wage-setting influences bargaining power of employees, and consequently the flexibility of real wages. Indicators of employment protection, on the other hand, reflect the extent of the labor market's inability to adjust to changes in the labor demand. Similarly, unemployment replacement rates affect the willingness of people to transition from unemployment into the workforce, resulting in a more rigid labor market. While our model below only captures these channels in a reduced form, most of the literature studies labor market imperfections in a closed-economy setting. A notable exception is by Bodenstein et al. (2018), who embed a search-and-matching labor market into a single-sector small

[10] For each country, we regress ULC on a constant and NER and collect the residuals, and add to them the country average ULC so as to preserve the correct cross-sectional information. Note that none of our results hinge on the use of either measure of the unit labor costs.

open economy model to study the effects of commodity price shocks. They find that the real exchange rate adjustment is a key force that induces a tightening of the labor market conditions observed in advanced commodity exporter countries.

The role of the supply-side labor wedges in causing real exchange rates has been studied in the eurozone (see Berka et al. 2018). To further this stream of the literature, we construct a panel of variables measuring differences in institutional labor markets relevant for the supply-side labor wedge. Although wedges likely also exist on the demand side of the labor market (e.g., due to product-market monopolies, see Karabarbounis 2014), the current paucity of usable data prohibits us from including these in our analysis. We use indicators from the Institutional Characteristics of Trade Unions, Wage Setting, State Intervention and Social Pacts (ICTWSS) database compiled by the OECD in conjunction with the Amsterdam Institute for Advanced Labor Studies (see Visser 2013 and https://www.oecd.org/employment/ictwss-database.htm). The institutional variables measure aspects of the labor markets pertaining to wage determination, while being largely orthogonal to productivity. For example, many characteristics of national wage bargaining structures have evolved over longer periods of history which make them largely exogenous at the medium-run frequencies we consider in our study. We choose summary variables that best capture the institutional differences in wage-setting while also having some time-series variability to be of practical use in our estimations (see Data Appendix sub-section **??** for details).

The ICTWSS database has been used in other macroeconomic literature, though not directly in studies of real exchange rates. Bertinelli et al. (2016) build a general equilibrium model of an open economy with a two-sector search-and-matching component for the labor market. In their model, wages differ between traded and nontraded sectors. Empirically, they find that wages in nontraded sector relative to traded sector decline following a shock to traded relative to nontraded TFP. This effect is stronger in countries with more regulated labor market, as measured by a variety of indicators in the ICTWSS database in their paper. Gnocchi et al. (2015) find that these indicators are related to cyclical movements in real wages, labor productivity and unemployment in OECD economies. Without attributing causality, Egert (2016) finds that anti-competitive regulations can in some cases be correlated with the total factor productivity measures, both in cross section and in time series across a panel of OECD countries [11].

[11] In an earlier version of this paper, with sample ending in 2007, we included two of the variables also used in Egert (2016): the strength of the employment protection laws and

Specifically, we consider the following labor market institutional variables from the OECD/AIAS ICTWSS database. Our preferred variable $COORD_{i,t}$ measures the coordination of wage bargaining by categorizing the strength of norms that guide wage-setting in a particular country, relative to Germany. In each country, categorical variable range from 1 (no coordination or fragmented coordination) to 5 (binding norms regarding maximum or minimum wages as a result of centralized bargaining or unilateral government imposition of wage schedule)[12]. We also utilize additional labour market institutional variables: $CENTRAL_{i,t}$, a summary index of the degree of centralization of wage bargaining, $CB_{i,t}$, measuring the right to central bargaining, $TUD_{i,t}$ the union density constructed using administrative and survey data, and $WB_{i,t}$, the predominant level at which wage bargaining takes place. We additional create a summary labor market indicator variable, either as the arithmetic average of the above, or as their principal component.

2.2 Description of prices and sectoral productivity

The Balassa-Samuelson hypothesis predicts a positive relationship between sectoral productivity differentials and the real exchange rate. Figure 7 shows the overall correlation between the RER and the relative TFP in positive (0.18), and that this is particularly true for the average levels of these variables. Countries with a higher relative productivity in the traded to nontraded sectors tend to have a higher level of q. Ireland and Belgium had the highest overall relative TFP levels over their individual samples relative to Germany, while Austria and Hungary the lowest. On the other hand, Sweden and

the unemployment benefit replacement ratio. Theoretically, both of these labor market frictions limit sectoral labor mobility, found to be relevant for explaining the dynamics of the relative nontraded to traded price in the OECD by Bertinelli et al. (2020). They were excluded from the current version because they significantly reduce the data sample.

[12] We choose this variable because it has a sufficient time variability within each country, while being comprehensive in its categorization of wage negotiation institutional framework as pertaining to the outcomes. The variable is coded as follows: 1 = Fragmented wage bargaining, confined largely to individual firms or plants, no coordination. 2 = Some coordination of wage setting, based on pattern setting by major companies, sectors, government wage policies in the public sector, etc. 3 = Procedural negotiation guidelines (recommendations on , for instance, wage demand formula relating to productivity or inflation). 4 = Non-binding norms and guidelines issued by the government, union and employers' associations, or extending from an extensive, regularized pattern setting coupled with high degree of union concentration and authority. 5 = Binding norms regarding maximum or minimum wage rates or wage increases issued as a result of a) centralized bargaining, with or without government involvement, or b) unilateral government imposition of wage schedules/freezes. The Codebook containing full variable descriptions can be found at https://www.oecd.org/els/emp/Codebook-OECD-AIAS-ICTWSS.pdf.

Denmark have the highest q levels, and the Czech Republic and Hungary the lowest. The correlation of annual growth rates of relative TFP and growth rates of RER is negative (-0.2), as shown in the third panel of Figure 7. This in line with the bulk of the time-series-based studies in the literature. Relative to Germany, Finland and Italy had the highest average growth rates of relative TFP over their respective samples, while the Netherlands and Hungary had the lowest. Similarly, the highest average rate of q appreciation rates were observed in the Czech Republic and Hungary, with the lowest in Italy and the US.

Finally, relative unit labor costs grew the most in Hungary and the Czech Republic, and fell the most in Japan and Austria (see Figure 2). The correlation between $OULC$ and q is 0.33 in the unbalanced panel.

3 Real Exchange Rates in a Theoretical Model

In this section we introduce a simple augmented Balassa-Samuelson model (see Balassa (1964) and Samuelson (1964)). The fundamental tenet of the Balassa-Samuelson mechanism is that higher productivity in traded sector at home raises wages in that sector and consequently in the whole economy. Facing rising wage costs without an increase in productivity, firms in the nontraded sector need to raise their prices, causing both an increase in the relative price of nontraded to traded goods, and in the overall price level. Thus, home real exchange rate appreciates. Conversely, a higher nontraded productivity depreciates home real exchange rate.

This mechanism has seen several key refinements over the decades. The introduction of a distribution sector effectively acts to magnify the Balassa-Samuelson mechanism (see Engel (1993), Froot and Rogoff (1995a), Engel and Rogers (1996) and Devereux (1999) for early expositions). The imperfect substitutability of traded goods limits the elasticity of the real exchange rate to traded productivity (Fitzgerald (2003) and Benigno and Thoenissen (2003)). Finally, limits to inter-sectoral labor mobility as well as other labor market restrictions open an avenue for labor market interference with real exchange rate determination that is orthogonal to productivity (see, e.g., Cardi and Restout (2015) and Bertinelli et al. (2020))[13]. In representative agent models,

[13] Balassa-Samuelson model structures tend to eliminate demand factors as drivers of real exchange rates. Models of demand-side permanent drivers of q typically assume either non-homothetic preferences, or a concentration of government consumption in nontraded sector (for example, Bhagwati (1984), Bergstrand (1991), and others).

labor wedges are used as a shortcut to capture labor market imperfections[14]. A supply-side labor wedge appreciates real exchange rate by raising real wages and unit labor costs, and consequently prices; a demand-side labor wedge appreciates real exchange rate by lowering demand for labor and real wages, while also increasing consumer prices.

We use a minor variation of the model in Berka et al. (2018). Theirs is a two-sector, two-country DSGE model with a distribution sector, imperfect elasticity of substitution in tradables, and an economy-wide supply-side labor wedge. Sectoral productivity and aggregate labor wedge shocks cause movements in real exchange rate. We extend the model by allowing for a possibility of a labor wedge that varies by sector. This is quite realistic, particularly for supply-side wedges. The sectoral differences in the union power are a plausible driver of a sector-specific labor wedges. Historically, collective wage bargaining has been performed at the levels of industries, and was one reason why union organizations used to be sector-specific. Differences in union movements between countries shaped the differences in wage negotiation practices and frameworks. Unionization rates still vary by sector within countries, at times dramatically (see OECD 1994 or OECD 1997). As an example, Figure 8 shows the unionization rates for the US traded and nontraded sectors.

Although the welfare consequences of fixed labor contracts were first pointed out by Leontief (1946), our current macroeconomic understanding of the roles played by the unions is largely based on the insider-outsider model. Lindbeck and Snower (1985) introduce the insider-outsider approach which vests some bargaining power to the employees ('insiders'), and discuss their implication for wage setting. Sollow (1985) adds a focus on skills and the longer-term relevance of the overall labor pool. In the first fully developed microeconomic treatment of the union's insider-outsider interaction, Lindbeck and Snower (1988) let the union insiders adopt a form of 'harassment' towards the non-union outsiders.[15] In equilibrium, insiders charge a wage which is a markup on the outside wage. We adopt this equilibrium result as an assumption in our representative agent model, and also assume that the outsiders' wage equals the marginal product of labor in a given industry. The union wages are then a markup on this marginal product.

[14]A labor wedge is defined as a gap between the marginal product of labor in production and the marginal rate of substitution between leisure and consumption of households. The literature points to both supply- and demand-side wedges arising from job search costs, income taxes, monopoly power in wage setting, nominal wage stickiness, etc.

[15]The insider-outsider approach has since been adopted chiefly to study employment (e.g., Blanchard and Summers 1986, and Lindbeck and Snower 2001), especially in Europe.

While the effects of labor unions on real exchange rates have been appreciated since Giovannini (1990), only a few studies propose a concrete mechanism to rationalize it. In a small open economy model with labor unions in nontraded sector, de Gregorio et al. (1994) study the relative price of nontraded to traded goods in Europe. In their model, the unions minimize a loss function $(L-\bar{L})^2+\sigma(w-\bar{w})^2$ where $\bar{L}$ and $\bar{w}$ are unions' targets for employment and real wage. In equilibrium, real exchange rates appreciate in real wage targets set by the unions.[16] Berka et al. (2018) show that when the labor wedge does not differ by sector, its effect on the real exchange rate is indistinguishable from a wedge that is modelled as parametric shifter of the disutility of labor.

We assume that the wage markup is as in Galí et al. (2007) and Karabarbounis (2014): $\mu_{j,t}=(w_t-p_{j,t})-MPL_{j,t},\ j\in\{T,N\}$, and similarly in the foreign country. The rest of the model is identical to the flexible-price version of Berka et al. (2018) and is explained in the Model Appendix B. To motivate our empirical specification, we focus on the solution of the linearized version of the model around a symmetric steady state when there is no home bias. Let q be the real exchange rate measured as the relative price of the home to foreign consumption basket, χ^R the relative (always home relative to foreign) disutility of labor used to proxy the homogeous labor wedge, a_T^R the relative productivity in the traded sector, a_N^R the relative productivity of the nontraded sector, μ_N^R the relative markup in the nontraded sector and $\mu_N^R-\mu_T^R$ the relative markup in the nontraded sector relative to traded sector. These relative markups allow for sector-specific labor wedges on the supply side. Then, real exchange rate q can be expressed as:

$$q=\alpha_\chi\chi^R+\alpha_T a_T^R+\alpha_N a_N^R+\alpha_{\mu_N}\mu_N^R+\alpha_{\mu_N-\mu_T}(\mu_N^R-\mu_T^R) \quad (1)$$

[16] An alternative model structure that would result in real wage markups can be akin to Ahn et al. (2017). Under the assumption that sectoral labor unions aggregate household labor supply in each sector, and that labor inputs have an elasticity of substitution that varies by sector (e.g. if supplying jobs to different occupations in a nontraded sector requires skills that are not as directly substitutable as those in a traded sector), union wages can be written as a sector-specific markup on the marginal costs:

$$\tilde{Y_t^T}=A_t^T L_t^T,\ \text{where } L_t^T=\left(\int_0^1 (L_{it}^T)^{\frac{\zeta^T-1}{\zeta^T}}di\right)^{\frac{\zeta^T}{\zeta^T-1}}$$

$$\tilde{W_t^T}=\frac{\zeta^T}{\zeta^T-1}MC_t^T$$

and similarly for nontraded sector. This gives rise to an industry-level wage that is a sector-specific markup on the marginal product of labor.

where

$$
\begin{aligned}
\alpha_\chi = \alpha_{\mu_N} &= \frac{\sigma(1-\gamma\kappa)}{B} \\
\alpha_{a_T} &= \frac{\sigma(1-\gamma\kappa)}{B}\gamma\kappa\psi(\kappa\lambda+\phi(1-\kappa)-1) \\
\alpha_{a_N} &= -\frac{\sigma(1-\gamma\kappa)}{B}[1+\psi(1+\gamma\kappa(\kappa\lambda+\phi(1-\kappa)-1))] \\
\alpha_{\mu_N-\mu_T} &= \frac{\sigma(1-\gamma\kappa)}{B}\gamma\kappa\psi(\kappa\lambda+\phi(1-\kappa))
\end{aligned}
$$

and

$$B = \sigma+\psi\left(1+\kappa\left[\sigma(\psi-\theta)+\gamma^2\kappa(1-2\sigma\theta)+\gamma(\sigma(\phi+2\theta+\kappa(\lambda-\phi-\psi+\theta))-2)\right]\right)$$

In a standard calibration[17], the coefficients in (1) are: $\alpha_\chi = \alpha_{\mu_N} = 0.22,\ \alpha_{a_T} = 0.26,\ \alpha_{a_N} = -0.71,\ \alpha_{(\mu_N-\mu_T)} = 0.33$.

Our model solution preserves the Balassa-Samuelson prediction that traded productivity typically appreciates q, while allowing this elasticity to become negative when the elasticity of substitution between Home and Foreign traded goods λ is low, as shown by Benigno and Thoenissen (2003). The solution also highlights the additional roles played by the relative disutility of labor, relative wage markup in nontraded sector, and the sectoral difference in wage markups. The effect of the relative nontraded wage markup on q is indistinguishable from the effect of the disutility of labor, while the gap between the nontraded and traded markups acts to further appreciate the real exchange rate.

The closed-form solution here isn't readily amenable to an empirical investigation because the disutility of labor is unobservable. We therefore use the approach outlined in Berka et al. (2018) and transform this solution into one using the observable unit labor costs. In a special case of our model with no distribution sector nor home bias, and when output is linear in labor, we can show that $q = (1-\gamma)(\tau + a_T^R - a_N^R + \mu_N^R - \mu_T^R)$ where τ is the endogeneous terms of trade. Defining unit labor costs as nominal wage divided by real output and expressing the wage difference using the first order conditions of the traded firm's labor decision $(w - w^* - s = \tau + a_T^R - \mu_T^R)$, we can express relative unit labor costs as $rulc = \tau + (1-\gamma)a_T^R - (1-\gamma)a_N^R - \mu_T^R$. This allows us to write the real exchange rate in this special case as:

$$q = (1-\gamma)rulc + \gamma(1-\gamma)a_T^R - \gamma(1-\gamma)a_N^R + (1-\gamma)\mu_N^R \tag{2}$$

[17] Specifically, when $\sigma = 2$, $\kappa = 0.6$ (so that the distribution sector accounts for 40% of retail tradable goods in equilibrium), $\theta = 0.7$, $\gamma = \omega = 0.5$, $\psi = 1$, $\phi = 0.25$ and $\lambda = 8$. We discuss these choices in the Appendix, and they are taken from Berka et al. (2018).

In this simplified version of the model, the disutility of labor parameter influences real exchange rate only through unit labor costs. This is also true in the general form of the model, but cannot be shown in a closed-form. In the empirical section which follows, we find that it is the institutional differences in wage negotiation regulations (that result in higher markups) which further appreciate real exchange rates beyond the direct effect of relative unit labor costs.[18]

4 Empirical Methodology

To estimate the model equation (2), we need measures of differences in the nontraded wage markups between countries. Alternatively, if we were to estimate a structural model solution equation (1), we would additionally need the relative wage markup of nontraded to traded sector, as well as a measure of the labor wedge (we construct a measure of the latter in section 5.1 below). Since we are unable to find data on sectoral wage markups, we instead use the summary measure of the level of coordination in wage bargaining $COORD_{i,t}$ described earlier. As we discussed in sections 2.1 and 3 in some detail, the economic literature commonly accepts, and models, the notion that higher unionization rates and more centralized or prescribed wage bargaining raise sectoral and overall wage markups by giving workers more bargaining power over the production surplus.

We estimate the empirical form of (2) using pooled OLS:

$$q_{i,t} = \alpha + \beta a_{T,i,t} + \gamma a_{N,i,t} + \delta oulc_{i,t} + \omega x_{i,t} + \epsilon_{i,t} \tag{3}$$

where $q_{i,t}$ is the bilateral RER of country i in year t, $a_{T,i,t}$ and $a_{N,i,t}$ are the bilateral differences in levels of traded and nontraded TFP, respectively, $oulc_{i,t}$ is the relative (orthogonalised) unit labor cost, and $x_{i,t}$ is a vector of variables describing institutional characteristics of country i's individual labor markets. All variables are bilateral relative to the US in our baseline specification. In section 5.2, we re-estimate our results with Germany as the base country. We also estimate equation (3) with fixed and random effects, which rely mainly on time-series variation to estimate slope coefficients[19]:

$$q_{i,t} = \alpha + \beta a_{T,i,t} + \gamma a_{N,i,t} + \delta oulc_{i,t} + \omega x_{i,t} + \eta_i + \epsilon_{i,t} \tag{4}$$

[18] Because we do not observe sector-specific institutional wage bargaining characteristics, we implicitly assume that countries with higher legislative restrictions on wage bargaining experience proportionally larger gaps between nontraded and traded sector wage markups.

[19] The fixed effect regressions allow for different constant intercepts, while the random effects estimation assumes that intercepts can vary across countries, but are assumed to be random variables.

where η_i are cross-sectional country effects. Finally, we include results from a cross-sectional regression which uses time-series average values for each country i from a balanced panel:

$$q_i = \alpha + \beta a_{T,i} + \gamma a_{N,i} + \delta oulc_i + \omega x_i + \epsilon_i \tag{5}$$

5 Empirical results

The benchmark results are summarised in Table 4.[20] We begin with a standard Balassa-Samuelson regression, where traded and nontraded TFP influence q with different magnitudes, and proceed by sequentially adding relative unit labor costs, and then indicators of labor market institutions.

In the pool regression, a_T is significant and with expected sign, while the elasticity for a_N is not significantly different from zero. A 1 percent improvement in traded TFP relative to Germany appreciates a country's q by 0.33 percent, ceteris paribus. In the fixed-effect and random-effect regressions of the basic Balassa-Samuelson model, relative nontraded TFP has the expected sign (-0.17 and -0.14, respectively), while the traded TFP is not significant.[21] Poor, and often counter-intuitive comovement of TFP and q comovement in time-series is a recurring result in the literature, especially for the OECD countries and in the traded sector.[22] In the cross-sectional regression, a_T is significant but a_N is not.

Adding our unit labor cost measure to the baseline regression model increases the significance of the parameter estimates (columns 5 to 8).[23] This is in line with the predictions of our model, in which relative ULC capture the homogenous effect of supply-side labor wedge, as seen in equation (2). Although nontraded TFP remains insignificant in the pool regression, but remains significant in the fixed- and random-effect regressions. Our results suggest that relative unit labor costs are particularly important in explaining the time-series movements of q that are unrelated to TFP. The regression fit

[20]Panel unit root tests reject non-stationarity over our benchmark sample, and do not reject the null of no cointegration for our default specification.

[21]Haussman tests indicate a preference for fixed over random effects specification.

[22]The literature finds more empirical support for the Balassa-Samuelson hypothesis in cross-section than in the time-series. When we lower frequency of observations by constructing 5-year non-overlapping averages of all our variables, we find that our baseline results are broadly qualitatively unchanged. We conclude our main results are not driven by higher-frequency movements in the data.

[23]We again note that these results cannot be driven by variation in nominal exchange rates because we orthogonalize relative ULC to such variation while constructing $OULC$. Even if we add NER to our regressions, $OULC$ remains highly significant.

improves, with $\bar{R}^2$ rising from 0.10 to 0.29 in the pool and from 0.76 to 0.88 in fixed-effect regressions.

Finally, we add our preferred measure of labor market institutional rigidity, $COORD_{i,t}$. The concentration of union membership at all levels is a significant additional driver in pool and cross-sectional regressions, in part due to a low frequency of changes in the institutional frameworks around wage bargaining (Columns 9 to 12). Labor markets with a higher degree of wage coordination tend to be associated with more appreciated real exchange rates, ceteris paribus. This is consistent with the role played by wage markups in our model. We conclude that, under the assumptions of our model, there is strong evidence to support the notion that real exchange rates are determined in a Balassa-Samuelson model augmented for labor wedge measures in our broad sample of OECD countries.

5.1 Labor wedge and real exchange rates

While labor wedge plays a key role in our model of real exchange rates, it is unobserved in reality. Like Berka et al. (2018), we use the model relationship between the unit labor costs, the labor wedge, and total factor productivity in both sectors to calculate the model-implied labor wedge in our data. We do this under the assumption of zero relative sectoral wage markup differences, which we have no data for.[24] We consequently re-estimate our empirical models with this labor wedge instead of $OULC$. Table 5 reports our findings.

Our results remain qualitatively unchanged when using the model-implied labor wedge calibration. For example, in the model with sectoral TFP and labor wedge, nontraded TFP becomes more significant, but traded TFP loses its significance in the fixed- and random-effect regressions[25]. This chnage in significance also holds in the augmented model, with an added institutional labor market variable $COORD$.

5.2 Role of the exchange rate flexibility

To add value to the literature which occasionally highlights the roles of exchange rate flexibility, we next consider the differences in our results between the nine eurozone members and nine countries with floating exchange

[24] That is, we calculate $lw_{i,t} = 0.33a_{T,i,t} + 2.33a_{N,i,t} + 2.8OULC_{i,t}$, as in Berka et al. (2018), because symmetric ($\mu_N^R = \mu_T^R$) flexible-price steady states are identical in these two models.

[25] Note that the model fit does not change, since we are not using any additional information in our regressions.

rates[26]. With Germany as the base country, we construct a dummy variable $euro_{i,t}$ which equals 1 if country i is a member of the eurozone in year t. This allows us to separate all bilateral country pairs into 'common currency' and 'floating exchange rate' groups. We consequently re-estimate our three models, with a model-implied labor wedge, allowing for the interaction slope dummies to capture the additional effects of the eurozone membership on the elasticities of our empirical model. We report these results in Table 6. Three points are worth noting here.

First, in the basic model, the time-series elasticity of a_T is significantly smaller in the eurozone than in floating exchange-rate countries. A plausible explanation is that the elasticity of substitution between traded goods in the eurozone is lower than for country pairs with floating exchange rates. The integration of supply chains amongst the eurozone member states that begun in 1950s allowed for a higher degree of specialization, particularly within the manufacturing sector (as exemplified by the German "Mittelstand") relative to non-eurozone country pairs. As has been documented elsewhere since at least Benigno and Thoenissen (2003) and Fitzgerald (2003), a lower elasticity of substitution between traded goods decreases the elasticity of a_T in the Balassa-Samuelson framework. This is exactly the result we observe here.

Second, these differences in the elasticity of a_T remain even after controlling for labor wedges. Additionally, our results with the labor wedge see a significantly higher elasticity of a_N in the Eurozone, and a significantly lower elasticity of the labor wedge LW in the Eurozone. A possible hypothesis that could explain the labor wedge result is that the eurozone's history of economic integration simultaneously resulted in deeper levels of specialization of the manufacturing industries, while also eroding the bargaining power of workers over time (note that the significance is primarily driven by the time series dimension of the data).

Finally, our augmented model adds the summary coordination of wage bargaining variable $COORD$. This inclusion significantly improves the goodness of fit of the regression in the cross-sectional dimension. The base-group effect of adding $COORD$ is positive and highly significant in all regressions, and its inclusion elucidates the Balassa-Samuelson mechanism in the base group[27]. As was the case with the labor wedge, the impact of the institutional differences in the levels of wage bargaining on real exchange rates is significantly lower in the eurozone. The inclusion of this institutional labour market variable

[26] Berka et al. (2018) solely focus on the eurozone.

[27] The inclusion of the $COORD$ variable increases the significance of the productivity coefficients.

dampens the differences in the elasticity of a_T between the eurozone and non-eurozone country pairs, but it does not eliminate them.

To summarise, we find significant differences of elasticities of traded TFP, in all models, of common-currency and floating exchange rate country pairs. The elasticity of a_T is lower in the eurozone, as could be explained by the history of economic integration. Some, but not all, of these differences can be explained by the labor wedge, and additionally by the institutional differences in the wage bargaining between countries. Labor wedge and $COORD$ themselves have a lower effect on the real exchange rates in the eurozone.

While it is beyond the scope of this paper to explore this finding in depth, we provide illustrative evidence about the role played by the trade integration in explaining the reported coefficient differences between common and floating exchange rate regimes.

5.3 Robustness

We consider three robustness checks for our results. First, in line with some older studies, we assume the absolute values of the elasticities for traded and nontraded TFP differentials are the same and estimate the baseline and the augmented models accordingly. The crux of our results carries through, as can be seen in Table 8. Relative (traded-to-nontraded) TFP $a_T - a_N$ is highly significant in pool and cross-section, and is significant in fixed- and random effects models if we also control for relative unit labor cost and labor market differences.

Second, we consider alternative measures of labor market institutions. Table 9 provides a summary of coefficient estimates across different labor market institutional variables, obtained mostly from the same data source. Many are significant when added to the benchmark model in a pooled regression. However, the only variables that are significant in both fixed- and random effect specifications are C_B, relative union density TUD, and the average of four indicators ($LMIaverage$). These results point to a variety of possible causal channels for labor market institutional differences to cause real exchange rates, a fertile area for future research.

Third, we consider the inclusion of terms of trade. Benigno and Thoenissen (2003) and Fitzgerald (2003) first showed that when countries produce different traded goods, real exchange rates are part-driven by an endogenous terms-of-trade effect which runs counter to the Balassa-Samuelson effect. The net effect then depends on the elasticity of substitution between home and foreign traded goods. Our model incorporates this possibility. In Table 10, we see that the addition of terms of trade to the benchmark model preserves the highly significant coefficient estimates in all specifications except two of the cross-sectional regressions.

Finally, we consider the role of demand-side factors of determining real exchange rates (for an overview, see Froot and Rogoff 1995b). With non-homothetic preferences an increase in demand appreciates real exchange rate (Bergstrand, 1991). De Gregorio et al. (1994), Chinn and Johnston (1997) and others show that a concentration of government expenditures in the nontraded sector gives a channel for the aggregate demand to influence the real exchange rate. Lane and Milesi-Ferretti (2002) focus on the role of net external wealth. We include long-run real interest rate differentials ($RIRDIFF$) to control for demand influence on real exchange rates. Table 10 also shows that the inclusion of an interest rate differential does not change our baseline results. In the pool regression, there is a slight change in the coefficient sizes but no change in their significance, while the $RIRDIFF$ is positive and significant. Qualitatively, these results carry through into fixed- and random-effect regressions.

6 Conclusion

We evaluate the augmented Balassa-Samuelson hypothesis using a newly constructed panel dataset of levels of sectoral TFP estimates and real exchange rate levels for 18 OECD countries between 1970 and 2020. We find that the Balassa-Samuelson mechanism is clearly visible after we control for differences in countries' labor market institutions and unit labor costs, reflecting the effects of country- and time-varying labor wedges. We augment the model in Berka et al. (2018) for sectoral differences in firms' wage markups as in Galí et al. (2007) and Karabarbounis (2014), and show that it implies a need for an additional measure of institutional labor market differences by sector, as well as a common labor wedge measure, to be added to an empirical framework to elucidate the Balassa-Samuelson relationship in the data. We find strong empirical support for this extension.

We also find that eurozone economies have significantly lower elasticities of real exchange rates to traded TFP than noneurozone economies. These differences can be partly accounted for by the inclusion of labor wedges in the empirical analysis, making the augmented Balassa-Samuelson model equally applicable to floating exchange rate and common currency OECD member states. We find that a key difference in explaining real exchange rates between these two country groups lies in a significantly lower elasticities of the institutional labor market characteristics in the eurozone.

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A Tables and Figures

Table 1: Summary statistics: average levels

Country	Sample	$\overline{a_T}$	$\overline{a_N}$	$\overline{a_T - a_N}$	$\overline{q}$	$\overline{oulc}$	$\overline{coord}$	$\overline{tot}$	$\overline{rirdiff}$
AUS	1982-2017	-0.14	0.02	-0.16	0.11	-0.41	-0.59	-0.05	-0.002
AUT	1980-2017	-0.48	-0.06	-0.42	-0.04	0.04	0.004	0.01	-0.003
BEL	1980-2017	0.19	0.06	0.12	0.04	-0.13	0.07	0.02	0.0001
CZE	1995-2017	-0.49	-0.26	-0.23	-0.57	-0.29	-1.28	0.02	-0.001
DNK	1980-2017	-0.17	0.13	-0.29	0.25	-0.24	0.04	0.02	0.012
ESP	1980-2017	-0.21	-0.02	-0.19	-0.18	-0.59	-0.38	0.02	0.002
FIN	1975-2017	-0.17	0.01	-0.18	0.22	-0.23	-0.15	0.01	0.009
FRA	1980-2017	-0.09	-0.07	-0.02	0.08	-0.17	-0.84	-0.01	-0.001
HUN	1995-2015	-0.64	-0.35	-0.29	-0.91	-0.32	-1.30	0.01	0.002
IRL	1988-2014	0.17	-0.12	0.28	0.06	-0.53	-0.36	0.02	0.005
ITA	1972-2017	-0.18	0.07	-0.25	-0.04	-0.54	-0.47	0.00	-0.005
JAP	1973-2015	-0.43	-0.37	-0.06	0.12	0.03	0.12	-0.01	-0.014
NLD	1979-2017	0.05	0.28	-0.22	0.03	-0.10	0.02	-0.01	-0.003
NZL	1982-2017	-0.04	0.18	-0.22	-0.08	-0.45	-0.95	-0.01	-0.009
SWE	1993-2016	-0.12	0.04	-0.16	0.29	-0.26	0.01	0.02	-0.006
UK	1972-2016	-0.09	-0.02	-0.06	0.15	-0.31	-0.56	-0.03	-0.008
USA	1970-2017	0.08	0.23	-0.15	-0.02	-0.22	-1.16	-0.03	-0.004

Each variable x is in logarithmic form (except real interest rates which are in levels), expressed as a bilateral difference of country i value minus the value for Germany. A $\overline{x}$ represents a time-series average. a_T is the traded TFP, a_N is the nontraded TFP, q is the real exchange rate, $oulc$ is the orthogonalised bilateral unit labor cost difference, $Coord$ is a measure of coordination of wage bargaining, expressed as the log difference relative to Germany, TOT is export over import price level relative to Germany, $RIRDIFF$ is the real long run interest rate differential with Germany. Sample dates refer to a_T.

Table 2: Summary statistics: time-series volatility

Country	Sample	$s(a_T)$	$s(a_N)$	$s(a_T\text{-}a_N)$	$s(q)$	$s(oulc)$	$s(coord)$	$s(tot)$	$s(rirdiff)$
AUS	1982-2017	0.14	0.045	0.143	0.149	0.160	0.422	0.026	0.029
AUT	1980-2017	0.05	0.033	0.075	0.083	0.071	0.032	0.014	0.009
BEL	1980-2017	0.07	0.097	0.069	0.042	0.093	0.251	0.014	0.019
CZE	1995-2017	0.09	0.035	0.083	0.264	0.143	0.370	0.029	0.012
DNK	1980-2017	0.07	0.037	0.047	0.062	0.110	0.139	0.010	0.019
ESP	1980-2017	0.08	0.150	0.107	0.116	0.106	0.294	0.020	0.021
FIN	1975-2017	0.18	0.061	0.181	0.101	0.130	0.263	0.016	0.016
FRA	1980-2017	0.04	0.038	0.064	0.033	0.075	0.300	0.014	0.016
HUN	1995-2015	0.08	0.039	0.056	0.292	0.165	0.236	0.046	0.030
IRL	1988-2014	0.09	0.054	0.066	0.109	0.152	0.663	0.012	0.029
ITA	1972-2017	0.10	0.162	0.158	0.087	0.120	0.467	0.022	0.036
JAP	1973-2015	0.09	0.080	0.103	0.200	0.159	0.112	0.049	0.030
NLD	1979-2017	0.07	0.048	0.079	0.040	0.086	0.139	0.022	0.014
NZL	1982-2017	0.05	0.040	0.067	0.145	0.193	0.527	0.032	0.042
SWE	1993-2016	0.04	0.028	0.044	0.147	0.080	0.179	0.012	0.021
UK	1972-2016	0.06	0.090	0.110	0.116	0.150	0.276	0.032	0.029
USA	1970-2017	0.08	0.074	0.103	0.150	0.133	0.396	0.023	0.023

$s(x)$ represents a the time-series standard deviation of variable x in country i (expressed as a difference of country i relative to Germany).

Table 3: Summary statistics: average annual growth rates

Country	Sample	$g(a_T)$	$g(a_N)$	$g(a_T\text{-}a_N)$	$g(q)$	$g(oulc)$
AUS	1982-2017	-0.01	-0.002	-0.01	-0.001	0.006
AUT	1980-2017	0.002	-0.001	0.003	0.006	0.003
BEL	1980-2017	0.0003	-0.008	0.008	0.001	0.003
CZE	1995-2017	0.009	0.001	0.008	0.03	0.017
DNK	1980-2017	-0.002	0.0001	-0.002	0.003	0.002
ESP	1980-2017	-0.004	-0.01	0.007	0.006	-0.0002
FIN	1975-2017	0.011	-0.002	0.013	-0.002	0.002
FRA	1980-2017	0.001	-0.002	0.003	-0.002	-0.0003
HUN	1995-2015	0.001	0.004	-0.003	0.011	-0.006
IRL	1988-2014	0.002	-0.003	0.003	0.001	-0.007
ITA	1972-2017	-0.004	-0.013	0.009	-0.003	-0.003
JAP	1973-2015	-0.002	-0.008	0.005	0.007	-0.002
NLD	1979-2017	-0.006	-0.001	-0.005	0.003	0.002
NZL	1982-2017	-0.002	-0.003	0.001	0.004	0.014
SWE	1993-2016	0.004	0.003	0.001	-0.01	-0.001
UK	1972-2016	-0.004	-0.007	0.003	-0.002	-0.002
USA	1970-2017	-0.003	-0.006	0.003	-0.003	0.002

$g(x)$ represents the average annual growth rate of variable x, expressed as $\mathrm{d}(\log(x))$. Each variable x in country i is relative to the German value. a_T is the Traded TFP, a_N is the nontraded TFP, q is the real exchange rate, *oulc* is the orthogonalised bilateral unit labor cost difference.

Figure 1: TFP levels (relative to Germany, log)

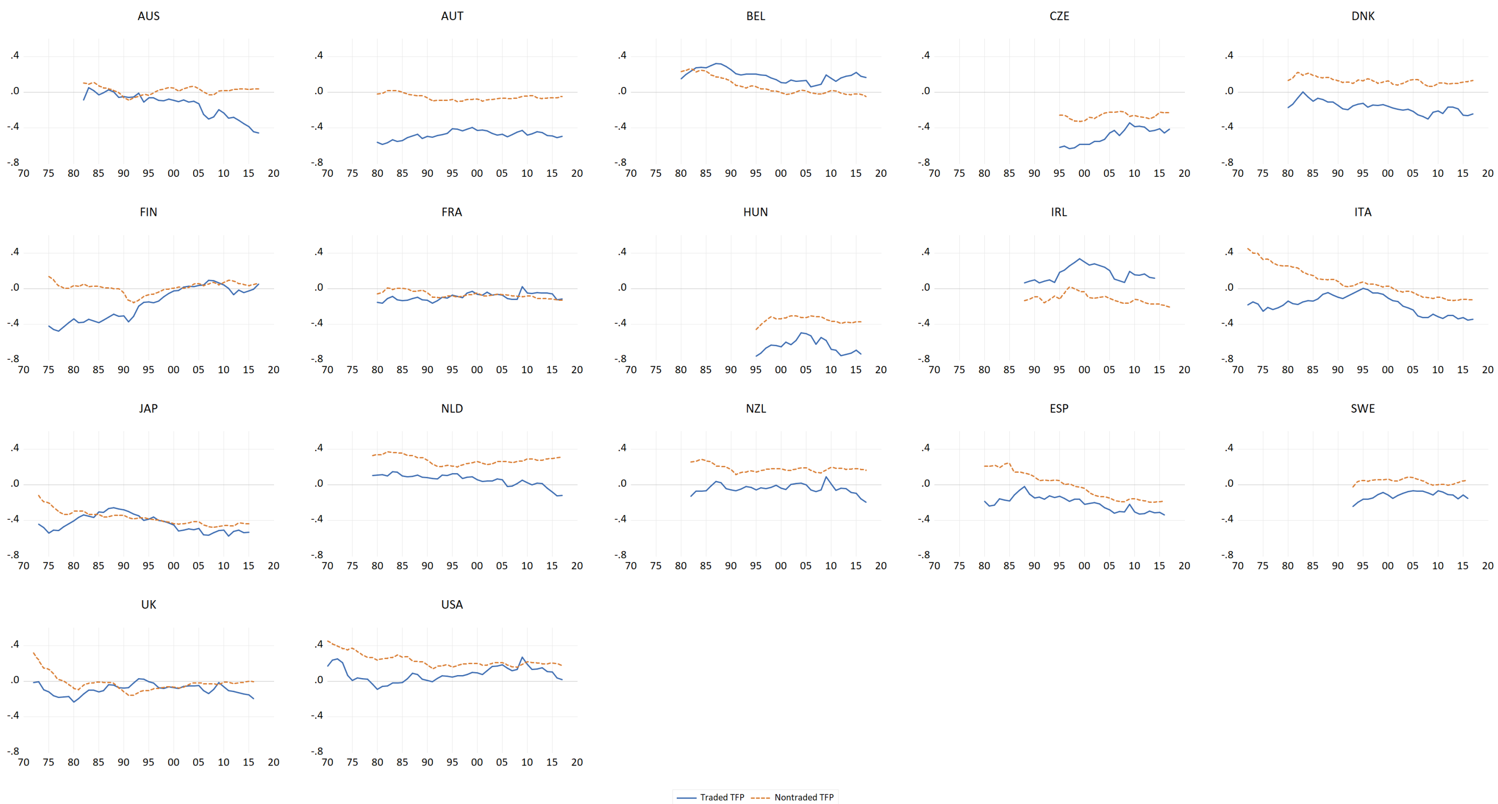


Figure 2: Levels of orthogonalized ULC ($OULC$) and ULC (relative to Germany, log)

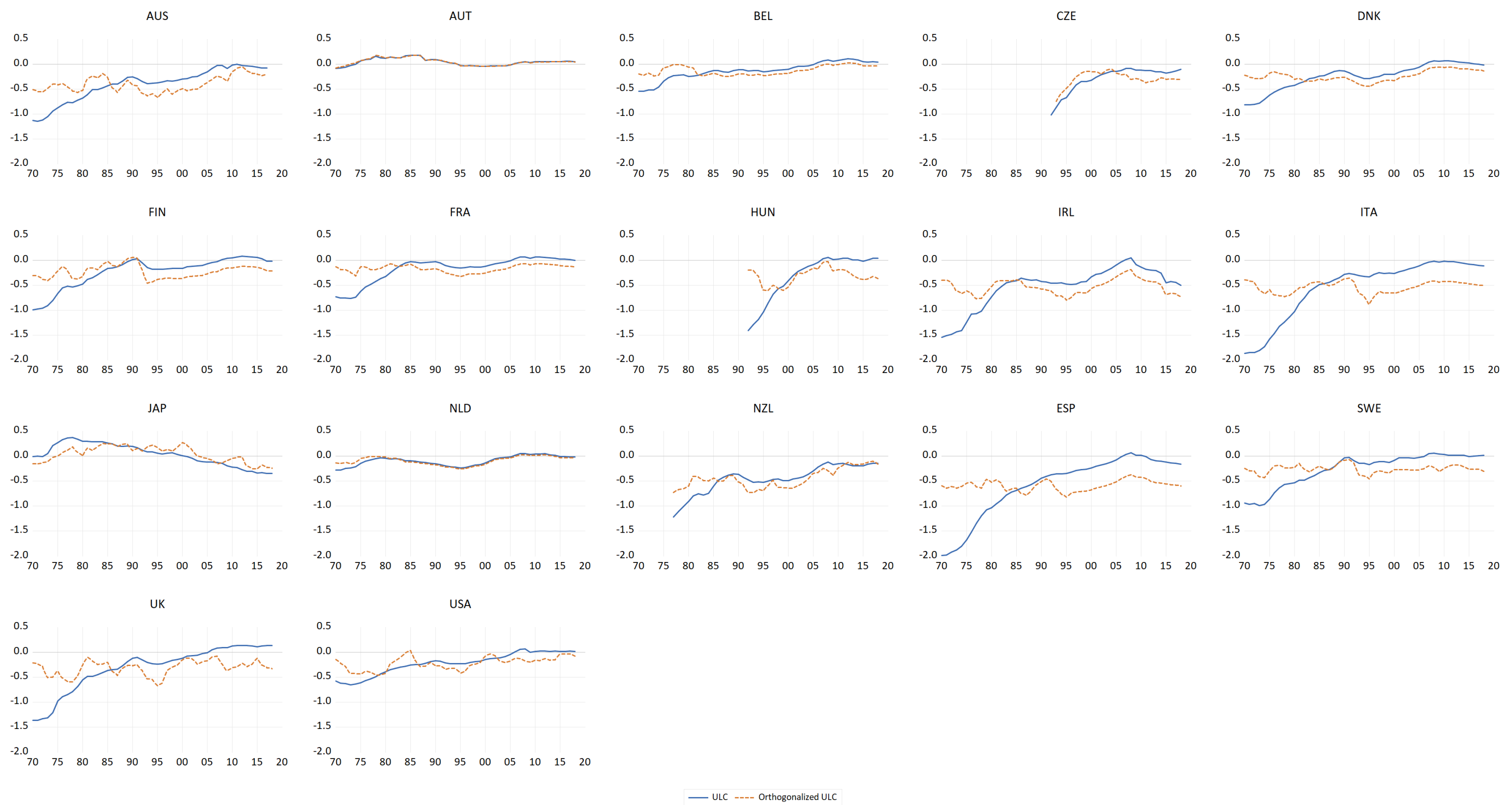
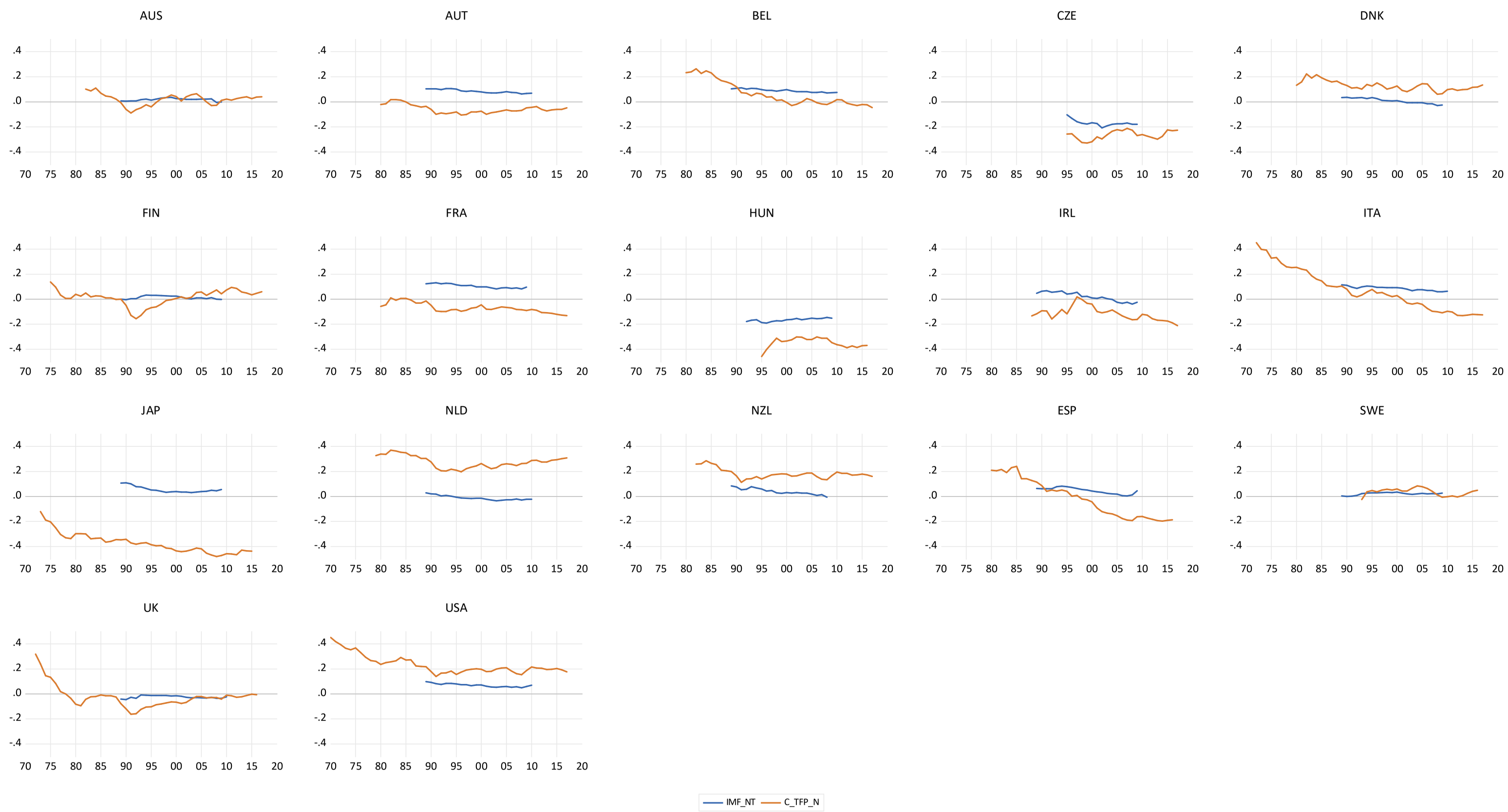


Figure 3: Levels of real exchange rates, orthogonalized ULCs and terms of trade (relative to Germany, logs)

Figure 4: Labor productivity vs TFP (Tradable levels, relative to Germany)

Source: Mano and Castillo (2015) and authors' calculations

Figure 5: Labor productivity vs TFP (Non-Tradable levels, relative to Germany)

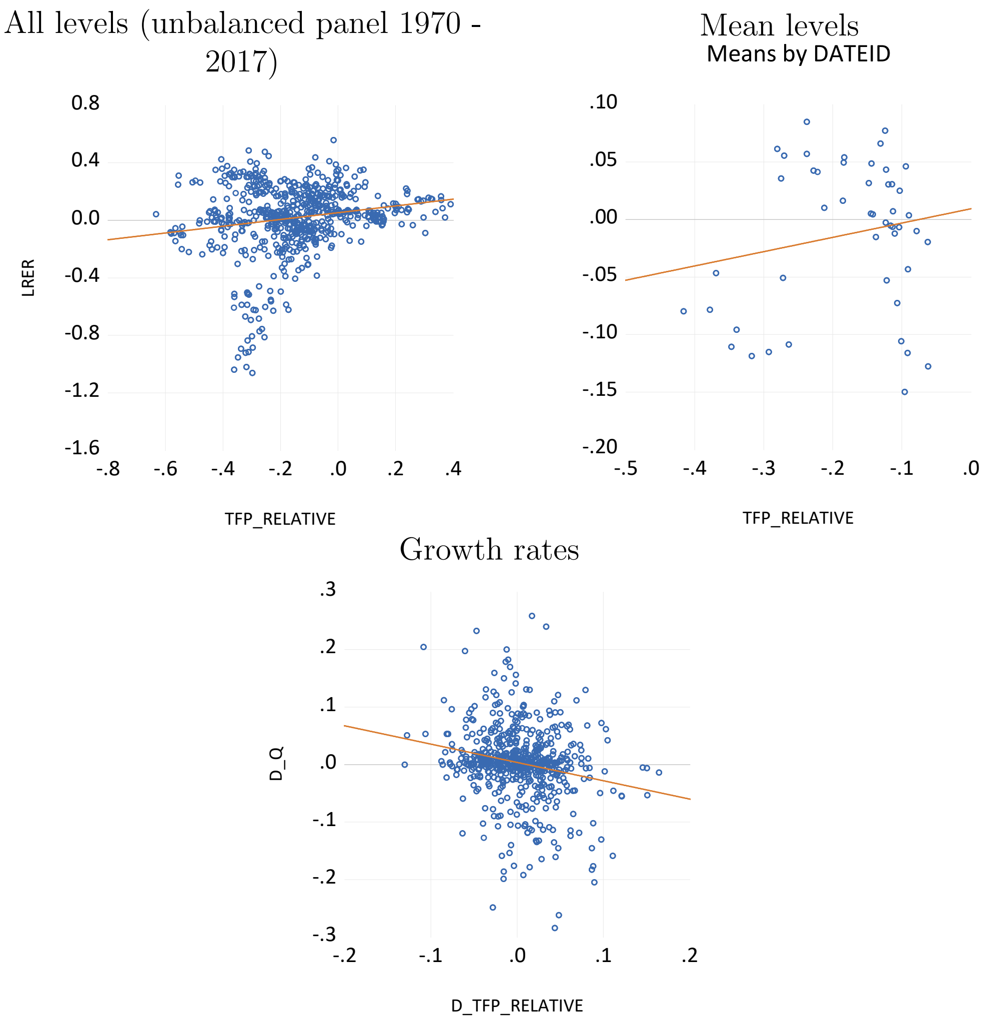


Source: Mano and Castillo (2015) and authors' calculations

Figure 6: Labor productivity vs TFP (Tradable-to-non-tradable levels, relative to Germany)

Source: Mano and Castillo (2015) and authors' calculations

Figure 7: Real exchange rate and cross-country productivity ratios



Note: All variables specified in log deviations from US levels. q is the bilateral real exchange rate in levels against the US based on aggregate CPI, a_T and a_N traded and nontraded TFP levels relative to the US. See Table **??** for country samples in the unbalanced panel.

Table 4: **** UPDATED 2 RER - TFP regressions (1970-2020)

	Basic model				Berka et al (2018) model				Augmented model			
	Pool	**FE**	**RE**	**XS**	**Pool**	**FE**	**RE**	**XS**	**Pool**	**FE**	**RE**	**XS**
a_T	0.33***	−0.10*	−0.07	0.61**	0.40***	0.08*	0.10**	0.70**	0.33***	0.08*	0.11**	0.48
s.e.	0.08	0.11	0.06	0.29	0.05	0.05	0.04	0.33	0.04	0.04	0.04	0.31
a_N	0.02	−0.17**	−0.14**	0.27	0.07	−0.11**	−0.09*	0.24	0.09*	−0.10**	−0.09*	0.22
s.e.	0.06	0.07	0.07	0.60	0.05	0.05	0.05	0.54	0.05	0.05	0.05	0.37
OULC					0.47***	0.63***	0.63***	0.49**	0.36***	0.63***	0.63***	0.10
s.e.					0.01	0.03	0.03	0.18	0.04	0.03	0.03	0.24
COORD									0.15***	-0.001	0.004	0.33***
s.e.									0.01	0.01	0.01	0.06
$\#obs$	621	621	621	17	621	621	621	17	621	621	621	17
$\bar{R}^2$	0.10	0.76	0.01	0.27	0.29	0.88	0.47	0.33	0.40	0.87	0.46	0.58
$Wald : \beta = -\gamma$	R***	R***	R***	N	R***	N	N	R*	R***	N	N	R*
LR	R***	R***	R***	R***	R***	N	N	R**	-	-	-	-
HT	-	-	R***	-	-	-	R**	-	-	-	R**	-

Dependant variable: q is log real exchange rate using aggregate CPI expressed as country i relative to the US. $a_{T,i,t}$ is an aggregation of 1-digit sectoral TFP of traded sectors using sectoral outputs as weights. $a_{N,i,t}$ is an aggregation of nontraded TFP. $OULC_{it}$ is orthogonalized relative unit labor costs calculated as are the residuals of a relative ULC regression on nominal exchange rate. $COORD$ is the coordination of wage-setting. 'Pool' is a pooled regression with all countries and periods sharing the same estimate of a constant and a slope. 'FE' is a fixed effect regression with countries as cross-sections. 'RE' is a random effects panel with countries as cross sections. 'XS' is a regression which uses the time-average value for each country and runs a cross sectional regression. Standard errors are in parentheses. The estimate of the constant is not reported. 'Wald' reports the result of the Wald test that $\beta = -\gamma$ (that coefficients on a_T and a_N are the same size but opposite signs, see equation (5)). The null for the likelihood ratio (LR) test is that the coefficient of $COORD$ is zero. 'R' denotes rejection of the null and 'N' non-rejection. All standard errors (except in cross section) are computed using panel-corrected standard errors method (Beck and Katz 1995) under the assumption of period correlation (cross-sectional clustering). The standard errors in cross section are Newey-West standard errors. A * denotes a 10%, ** 5% and *** 1% significance.

Table 5: **** UPDATED 2 RER - TFP regressions with Labor Wedge (1970-2020)

	Basic model				Berka et al (2018) model				Augmented model			
	Pool	**FE**	**RE**	**XS**	**Pool**	**FE**	**RE**	**XS**	**Pool**	**FE**	**RE**	**XS**
a_T	0.33***	−0.10*	−0.07	0.61**	0.35***	0.01	0.02	0.63*	0.29***	0.01	0.03	0.47
s.e.	0.08	0.11	0.06	0.29	0.05	0.04	0.04	0.30	0.04	0.04	0.04	0.28
a_N	0.02	−0.17**	−0.14**	0.27	−0.32***	−0.63***	−0.61***	-0.15	−0.21***	−0.63***	−0.61***	0.13
s.e.	0.06	0.07	0.07	0.60	0.06	0.06	0.06	0.56	0.06	0.06	0.06	0.42
LW					0.17***	0.22***	0.22***	0.18**	0.13***	0.22***	0.22***	0.04
s.e.					0.01	0.06	0.01	0.07	0.01	0.01	0.01	0.08
COORD									0.15***	-0.001	0.004	0.33***
s.e.									0.01	0.01	0.01	0.05
$\#obs$	621	621	621	17	621	621	621	17	621	621	621	17
$\bar{R}^2$	0.10	0.76	0.01	0.27	0.29	0.88	0.47	0.33	0.40	0.87	0.46	0.58
$Wald: \beta = -\gamma$	R***	R***	R***	N	N	R***	R***	N	N	R***	R***	R**
LR	R***	R***	R***	R***	R***	N	-	R**	-	-	-	-
HT	-	-	R***	-	-	-	R***	-	-	-	R***	-

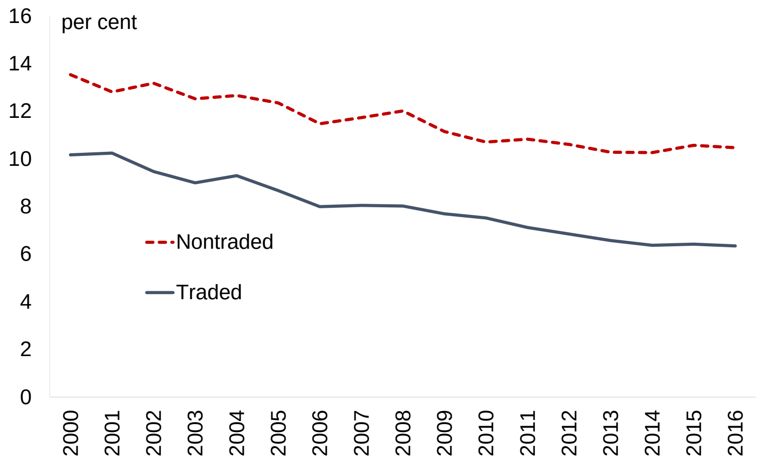
Dependant variable: q is log real exchange rate using aggregate CPI expressed as country i relative to the US. $a_{T,i,t}$ is an aggregation of 1-digit sectoral TFP of traded sectors using sectoral outputs as weights. $a_{N,i,t}$ is an aggregation of nontraded TFP. $OULC_{it}$ is orthogonalized relative unit labor costs calculated as are the residuals of a relative ULC regression on nominal exchange rate. $COORD$ is the coordination of wage-setting. 'Pool' is a pooled regression with all countries and periods sharing the same estimate of a constant and a slope. 'FE' is a fixed effect regression with countries as cross-sections. 'RE' is a random effects panel with countries as cross sections. 'XS' is a regression which uses the time-average value for each country and runs a cross sectional regression. Standard errors are in parentheses. The estimate of the constant is not reported. 'Wald' reports the result of the Wald test that $\beta = -\gamma$ (that coefficients on a_T and a_N are the same size but opposite signs, see equation (5)). The null for the likelihood ratio (LR) test is that the coefficient of $COORD$ is zero. 'R' denotes rejection of the null and 'N' non-rejection. All standard errors (except in cross section) are computed using panel-corrected standard errors method (Beck and Katz 1995) under the assumption of period correlation (cross-sectional clustering). The standard errors in cross section are Newey-West standard errors. A * denotes a 10%, ** 5% and *** 1% significance.

Table 6: **** UPDATED 2 **** RER - TFP regressions with Germany as base country, Labor Wedge, and Euro fixed effects (1970 - 2020)

	Basic model				**Berka et al (2018) model**				**Augmented model**			
	Pool	**FE**	**RE**	**XS**	**Pool**	**FE**	**RE**	**XS**	**Pool**	**FE**	**RE**	**XS**
a_T	0.59***	0.08	0.14*	1.04*	0.39***	0.13**	0.15**	1.06	0.21***	0.11**	0.17***	0.96**
s.e.	0.01	0.08	0.05	0.52	0.08	0.06	0.06	1.41	0.07	0.06	0.06	0.42
a_N	-0.09	-0.17	-0.1	−0.12	−0.28***	−0.79***	−0.74***	-0.42	0.08	−0.90***	−0.77***	0.30
s.e.	0.08	0.15	0.14	0.87	0.08	0.11	0.11	1.20	0.08	0.12	0.11	0.43
LW					0.22***	0.26***	0.26***	0.16	0.13***	0.26***	0.26***	−0.20
s.e.					0.02	0.01	0.01	0.32	0.02	0.01	0.01	0.15
COORD									0.21***	0.04**	0.04**	0.47***
s.e.									0.02	0.02	0.02	0.10
a_T*euro	−0.45***	−0.34***	−0.38***	−0.74	−0.19**	−0.27***	−0.28***	−0.72	0.01	−0.26***	−0.29***	-0.65
s.e.	0.08	0.11	0.11	0.53	0.09	0.08	0.08	1.34	0.08	0.08	0.08	0.45
a_N*euro	-0.14	0.03	-0.04	-0.11	−0.40***	0.35**	0.31**	-0.55	−0.84***	0.46***	0.32**	−1.17***
s.e.	0.11	0.17	0.16	0.78	0.10	0.13	0.13	1.15	0.10	0.14	0.13	0.30
LW*euro					−0.06***	−0.1***	−0.1***	0.04	0.05**	−0.11***	−0.10***	0.38**
s.e.					0.02	0.02	0.02	0.29	0.02	0.02	0.02	0.12
COORD*euro									−0.23***	−0.07***	−0.06**	-0.43
s.e.									0.03	0.03	0.02	0.26
$\#obs$	621	621	621	17	621	621	621	17	621	621	621	17
$\bar{R}^2$	0.16	0.77	0.03	0.26	0.38	0.88	0.50	0.28	0.53	0.88	0.49	0.70
$Wald: \beta = -\gamma$	R***	N	N	N	R*	R***	R***	N	R***	R***	R***	R**
LR	R***	R*	N	N	R***	N	N	R***	-	-	-	-
HT	-	-	R**	-	-	-	N	-	-	-	R***	-

Dependant variable: q is log real exchange rate using aggregate CPI expressed as country i relative to the US. $a_{T,i,t}$ is an aggregation of 1-digit sectoral TFP of traded sectors using sectoral outputs as weights. $a_{N,i,t}$ is an aggregation of nontraded TFP. $OULC_{it}$ is orthogonalized relative unit labor costs calculated as are the residuals of a relative ULC regression on nominal exchange rate. $CONC$ is the concentration of union membership (weighting of sectoral and aggregate). 'Pool' is a pooled regression with all countries and periods sharing the same estimate of a constant and a slope. 'FE' is a fixed effect regression with countries as cross-sections. 'RE' is a random effects panel with countries as cross sections. 'XS' is a regression which uses the time-average value for each country and runs a cross sectional regression. Standard errors are in parentheses. The estimate of the constant is not reported. 'Wald' reports the result of the Wald test that $\beta = -\gamma$ (that coefficients on a_T and a_N are the same size but opposite signs, see equation (5)). The null for the likelihood ratio (LR) test is that the coefficient of $COORD$ is zero. 'R' denotes rejection of the null and 'N' non-rejection. All standard errors (except in cross section) are computed using panel-corrected standard errors method (Beck and Katz 1995) under the assumption of period correlation (cross-sectional clustering). The standard errors in cross section are Newey-West standard errors. A * denotes a 10%, ** 5% and *** 1% significance.

Figure 8: Traded and nontraded average unionization rates in the US



Source: BLS `https://www.bls.gov/webapps/legacy/cpslutab3.htm`

Table 7: Average unexplained real exchange rate levels

	Basic model	**Augmented model**	**Unconditional** q
AUS	0.06	0.21	0.02
AUT	0.17	0.10	0.02
BEL	0.14	0.06	-0.06
CZE	-0.54	-0.49	-0.81
DNK	0.32	0.44	0.31
ESP	-0.02	-0.03	-0.14
FIN	0.31	0.24	0.21
FRA	0.23	0.17	0.09
GER	0.11	-0.02	0.00
HUN	-0.74	-0.60	-0.89
IRE	0.14	0.18	0.11
ITA	0.00	0.10	-0.03
JPN	0.61	0.29	0.25
NLD	-0.03	0.13	-0.04
NZL	-0.25	0.15	-0.18
SWE	0.26	0.31	0.26
UK	0.37	0.09	0.22
Average (absolute)	0.25	0.21	0.21

The figure reports total fixed effect estimates from the benchmark specification in Table 4 for the sample 1990-2007. Each number represents the sum of the constant and the fixed effect estimates for a given country.

Table 8: *** UPDATED *** Robustness to use of relative TFP measure (1970-2020)

	Pool			Fixed effects			Random effects			Cross-sectional		
	1	**2**	**3**	**4**	**5**	**6**	**7**	**8**	**9**	**10**	**11**	**12**
$a_T - a_N$	0.23***	0.21***	0.27***	0.01	0.09**	0.22***	0.02	0.10**	0.23***	0.51*	0.34	0.45
s.e.	0.05	0.04	0.04	0.08	0.04	0.05	0.05	0.04	0.05	0.25	0.26	0.27
OULC		0.28***			0.63***			0.63***			−0.08	
s.e.		0.04			0.03			0.03			0.26	
LW			0.14***			0.19***			0.19***			0.15
s.e.			0.01			0.01			0.01			0.09
COORD		0.16***	0.15***		−0.002	−0.02		0.003	−0.01		0.41***	0.31***
s.e.		0.01	0.01		0.013	0.015		0.01	0.01		0.15	0.09
# **obs**	621	621	621	621	621	621	621	621	621	17	17	17
$\bar{R}^2$	0.03	0.30	0.40	0.76	0.88	0.85	-0.001	0.47	0.36	0.02	0.41	0.51

Dependant variable: q is log real exchange rate using aggregate CPI expressed as country i relative to the US. a_i is the log of TFP level of traded relative to nontraded sector in country i ($a_{T,i,t} - a_{N,i,t}$) relative to the US. $OULC_{it}$ is orthogonalized relative unit labor costs calculated as are the residuals of a relative ULC regression on nominal exchange rate (expressed at the correct average level). x proxied using $CONC$, defined as the centralization of wage bargaining (weighting of sectoral and aggregate), specified as up for a more centralised labor market. The estimate of the constant is not reported.

Table 9: **** UPDATED **** Coefficient estimates of selected labor indicators in the benchmark specification (1970 - 2020)

	CENTRAL	C_B	TUD	W_B	W_S	LMI average	LMI PC
Pool	0.05**	0.02	0.12***	0.05**	0.15***	0.14**	0.03***
FE	-0.02	-0.03**	-0.18**	-0.01	-0.001	-0.04**	-0.01
RE	-0.02	-0.03**	-0.15**	-0.01	0.004	-0.04**	-0.005
XS	0.12	0.05	0.19***	0.12	0.33***	0.29**	0.06**

A * denotes a 10%, ** 5% and *** 1% significance when one labor market indicator is added to the benchmark specification from Table 4. 'Pool' is a pooled regression with all countries and periods sharing the same estimate of a constant and a slope. 'FE' is a fixed effect regression with countries as cross-sections. 'RE' is a random effects panel with countries as cross sections. 'XS' is a regression which uses the time-average value for each country and runs a cross sectional regression. Standard errors are in parentheses.

Table 10: **** UPDATED **** Robustness of RER-TFP regressions to adding terms of trade

	Pool			FE			RE			XS		
	1	**2**	**3**	**1**	**2**	**3**	**1**	**2**	**3**	**1**	**2**	**3**
$a_T - a_N$	0.21***			0.16***			0.17***			0.39		
s.e.	0.04			0.05			0.04			0.27		
a_T		0.21***	0.24***		-0.03	−0.20***		0.002	−0.16***		0.41	0.61*
s.e.		0.04	0.05		0.04	0.05		0.04	0.05		0.29	0.33
a_N		−0.21***	−0.07		−0.54***	-0.002		−0.52***	0.01		0.16	0.26
s.e.		0.05	0.05		0.06	0.07		0.06	0.84		0.32	0.61
LW	0.11***	0.11***		0.18***	0.18***		0.18***	0.21***		0.09	-0.012	
s.e.	0.01	0.01		0.01	0.01		0.01	0.01		0.09	0.08	
COORD	0.14***	0.14***		0.005	0.005		0.01	0.01		0.39***	0.41***	
s.e.	0.01	0.01		0.01	0.01		0.01	0.01		0.09	0.04	
TOT	−1.42**	−1.44**	−1.32***	-0.001	-0.36	0.44	−0.05	−0.38*	0.38	−4.73**	−3.9**	-0.50
s.e.	0.22	0.22	0.26	0.25	0.24	0.31	0.23	0.23	0.29	2.25	1.61	2.97
RIR	1.4***	1.9***	1.4***	1.39***	0.93***	1.37***	1.40***	0.98***	1.40***	4.0	0.30	-3.22
s.e.	0.19	0.19	0.3	0.2	0.19	0.24	0.19	0.19	0.23	9.2	9.1	15.5
# obs	599	599	599	599	599	599	599	599	599	17	17	17
R2	0.43	0.43	0.12	0.87	0.87	0.76	0.43	0.51	0.09	0.53	0.59	0.16

Dependant variable: q is log real exchange rate using aggregate CPI expressed as country i relative to the US. a_i is the log of TFP level of traded relative to nontraded sector in country i $(a_{T,i,t} - a_{N,i,t})$ relative to the US. $a_{T,i,t}$ is an aggregation of 1-digit sectoral TFP of traded sectors using sectoral outputs as weights. $a_{N,i,t}$ is a TFP aggregation of nontraded sectors. $OULC_{it}$ is orthogonalized relative unit labor costs calculated as are the residuals of a relative ULC regression on nominal exchange rate (expressed at the correct average level). TOT is export over import price *levels* expressed in logs relative to the US. The data sample is 1990-2007 (see Table **??**). 'Pool' is a pooled regression with all countries and periods sharing the same estimate of a constant and a slope. 'Fixed effects' is a panel regression with countries as cross-sections. 'Random effects' is a random effects panel with countries as cross sections. 'Cross-section' is a regression which uses the time-average value for each country and runs a cross sectional regression. Standard errors are in parentheses. The estimate of the constant is not reported. A * denotes a 10%, ** 5% and *** 1% significance.

B Model Appendix

This appendix section describes the model, focusing on the material added to the model of Berka et al. (2018). There are two countries, each populated by an infinitely-lived representative agent maximizing:

$$U_t = E_0 \sum_{t=0}^{\infty} \beta^t \left(\frac{C_t^{1-\sigma}}{1-\sigma} - \chi_t \frac{N_t^{1+\psi}}{1+\psi} \right), \quad \beta < 1. \tag{6}$$

where C_t is a composite consumption bundle and N_t is the supply of labor, and χ is a country-specific time-varying disutility of labor supply. The composite consumption good is a CES aggregator of traded and nontraded composite consumption (C_T and C_N). Traded consumption is a composite of home or foreign traded consumption goods (C_H and C_F). In line with the literature, these traded consumption goods at the retail level are CES aggregates of pure wholesale traded product and a retail input V which is nontraded. Hence, at home:

$$\begin{aligned}
C_t &= \left(\gamma^{\frac{1}{\theta}} C_{Tt}^{1-\frac{1}{\theta}} + (1-\gamma)^{\frac{1}{\theta}} C_{Nt}^{1-\frac{1}{\theta}} \right)^{\frac{\theta}{\theta-1}} \\
C_{Tt} &= \left(\omega^{\frac{1}{\lambda}} C_{Ht}^{1-\frac{1}{\lambda}} + (1-\omega)^{\frac{1}{\lambda}} C_{Ft}^{1-\frac{1}{\lambda}} \right)^{\frac{\lambda}{\lambda-1}} \\
C_{Ht} &= \left(\kappa^{\frac{1}{\phi}} I_{Ht}^{1-\frac{1}{\phi}} + (1-\kappa)^{\frac{1}{\phi}} V_{Ht}^{1-\frac{1}{\phi}} \right)^{\frac{\phi}{\phi-1}} \\
C_{Ft} &= \left(\kappa^{\frac{1}{\phi}} I_{Ft}^{(1-\frac{1}{\phi}} + (1-\kappa)^{\frac{1}{\phi}} V_{Ft}^{1-\frac{1}{\phi}} \right)^{\frac{\phi}{\phi-1}}
\end{aligned}$$

In the above equations, θ, λ and ϕ are elasticities of substitution between traded and nontraded goods, home and foreign tradables, and the wholesale traded good and nontraded input in retail sectors. γ, ω and κ are the steady-state shares of traded consumption in overall consumption, home bias in traded goods, and the weight of wholesale consumption in overall traded retail bundle. The optimal price indexes are:

$$\begin{aligned}
P_t &= \left(\gamma P_{Tt}^{1-\theta} + (1-\gamma) P_{Nt}^{1-\theta} \right)^{\frac{1}{1-\theta}}, \\
P_{Tt} &= \left(\omega \tilde{P}_{Ht}^{1-\lambda} + (1-\omega) \tilde{P}_{Ft}^{1-\lambda} \right)^{\frac{1}{1-\lambda}}, \\
\tilde{P}_{Ht} &= \left(\kappa P_{Ht}^{1-\phi} + (1-\kappa) P_{Nt}^{1-\phi} \right)^{\frac{1}{1-\phi}} \\
\tilde{P}_{F} &= \left(\kappa P_{Ft}^{1-\phi} + (1-\kappa) P_{Nt}^{1-\phi} \right)^{\frac{1}{1-\phi}}
\end{aligned}$$

where P_T and P_N are home country's price indexes of traded and nontraded aggregates, $\tilde{P}_H$ and $\tilde{P}_F$ are price indexes of Home and Foreign retail traded goods, and P_H and P_F are prices of Home and Foreign wholesale traded goods, measured at Home. We assume that law of one price holds in traded goods at wholesale level, and so $SP_H = P_H^*$ and $SP_F = P_F^*$. The real exchange rate is defined as

$$Q_t = \frac{P_t S}{P_t^*}$$

In our world of complete risk sharing, marginal utilities of consumption must equal between countries, when expressed in the same currency:

$$\frac{C_t^{-\sigma}}{P_t} = \frac{C_t^{*-\sigma}}{P_t^*} \tag{7}$$

The first order conditions imply the usual sets of equations. The implicit labor supply is governed by:

$$W_t = \chi_t P_t C^\sigma N_t^\psi$$

Where W_t is the nominal wage. The demand equations for consumption components are given by:

$$C_{Tt} = \gamma \left(\frac{P_{Tt}}{P_t}\right)^{-\theta} C_t, \qquad C_{Nt} = (1-\gamma)\left(\tfrac{P_{Nt}}{P_t}\right)^{-\theta} C_t$$

$$C_{Ht} = \omega \left(\frac{\tilde{P}_{Ht}}{P_{Tt}}\right)^{-\lambda} C_{Tt}, \qquad C_{Ft} = (1-\omega)\left(\tfrac{\tilde{P}_{Ft}}{P_{Tt}}\right)^{-\lambda} C_{Tt}$$

$$I_{Ht} = \kappa\omega \left(\frac{P_{Ht}}{\tilde{P}_{Ht}}\right)^{-\phi} \left(\frac{\tilde{P}_{Ht}}{P_{Tt}}\right)^{-\lambda} C_{Tt}, \quad I_{Ft} = \kappa(1-\omega)\left(\tfrac{P_{Ft}}{\tilde{P}_{Ft}}\right)^{-\phi}\left(\tfrac{\tilde{P}_{Ft}}{P_{Tt}}\right)^{-\lambda} C_{Tt}$$

Foreign consumption bundles, foreign prices, and demand first order conditions, are determined in an analogous fashion, and denoted with an *. Firms in each sector produce using labor and a fixed capital stock: $Y_{Nt} = A_{Nt} N_{Nt}^\alpha$, $Y_{Ht} = A_{Tt} N_{Ht}^\alpha$.

As described earlier, we allow for the existence of sectoral firms-side labor wedges, which can be motivated by the existence of sectoral labor unions. Specifically, we model them as sector-specific price markups $\mu_i,\ i \in (T, N)$ exactly as in Galí et al. (2007) and Karabarbounis (2014):

$$\mu_{j,t} = p_{j,t} - (w_t - MPL_{j,t}),\ j \in \{T, N\}$$

Ceteris paribus, μ raises firm's prices and appreciates q. When $\mu_T \neq \mu_N$, there is an additional effect of the differential sectoral labor wedge.

There are many papers that feature a wedge between the marginal rate of substitution in consumption and the marginal product in production. This literature is largely focused on understanding how labor market inefficiencies might affect labor supply. Sources of a 'labor wedge' could include many factors, including search costs, monopoly power in wage-setting, or sticky nominal wages (see Hall 1997, Chari et al. 2002, Galí et al. 2007, Shimer 2009, Karabarbounis 2014).[28] Irrespective of the underlying source of the wedge, these translate into price changes that are independent of TFP.[29]

We assume that prices are flexible and firms engage in monopolistic competition that yields the usual markup-pricing rule. Monetary policy in each country is characterized by a Taylor-type rule which adjusts nominal interest rates at home as follows:

$$r_t = \rho + \sigma_p \pi_t + \sigma_q (q_t - u_t)$$

where σ_p and σ_q are weights on inflation and real exchange rate stability, respectively, and u_t is a monetary policy shock (see Steinsson 2008). A similar monetary policy rule is followed by a foreign country. It can be shown that this implies that the nominal exchange rate in a symmetric equilibrium is a linear function of the differential monetary policy shocks $s_t = x(u_t^* - u_t)$ where x is a constant.

We focus here on the role of firm-side labor wedges, both between sectors and between countries, in driving the real exchange rate dynamics, in addition to Berka et al. (2018). The Ballassa-Samuelson mechanism implies that sectoral productivity differences influence real exchange rates. An increase in the Home relative (traded vs. nontraded) productivity over the Foreign appreciates the Home real exchange rate. An additional mechanism exists in models where traded goods are imperfect substitutes (such as here): increases in traded productivity additionally lowers the price of home exportables, thus depreciating the terms of trade and the real exchange rate. In usual

[28] Benassy-Quere and Coulibaly (2014) add product-market markups to the model of De Gregorio et al. (1994) and show empirically that if markups reflect product market regulations and employment protection, these have a meaningful impact on the eurozone's real exchange rates.

[29] Hall (1988) and Hall (1989) show that imperfect competition implies that measured TFP will itself be affected by demand fluctuations. One way to address this criticism would be to explicitly include estimates of markups for tradables and non-tradables, which is empirically infeasible as far as we are aware.

model calibrations, as well as in empirical studies, the former effect dominates the latter, and relative technological improvements are associated with real exchange rate appreciations.

At the core of both of these mechanisms lies the assumption that labor markets are perfectly competitive, and factors of production receive their marginal products. But there are clear differences in the efficiency of labor market institutions over time (owing to reforms) and also between countries. Such institutional differences play a prominent role in the assessment of international competitiveness. The traded sector first order conditions imply that an international wage difference can be decomposed into endogenous terms of trade movements, productivity differences, and markup differences:

$$w + s - w^* = \tau + a_T - a_T^* - (\mu_T^* - \mu_T)$$

where $\tau \equiv p_H - p_F^* - s$ is the terms of trade. A similar condition can be expressed using the nontraded sectors' first order conditions. With intranational labor market integration, wages equalise between sectors, which consequently implies that:

$$p_N + s - p_N^* = \tau + [a_T - a_T^* - (a_N - a_N^*)] + [\mu_N - \mu_N^* - (\mu_T - \mu_T^*)]$$

Thus, the real exchange rate for nontraded goods is a function of terms of trade, relative productivities (the Balassa-Samuelson effect) and relative markup differences. If we further assumed that $\kappa = 1$ and $\omega = 0.5$, so that the retail sector does not use nontraded inputs and there is no home bias in traded consumption, we could rewrite the above condition as:

$$p_n = [a_T - a_T^* - (a_N - a_N^*)] + [\mu_N - \mu_N^* - (\mu_T - \mu_T^*)]$$

where $p_n \equiv p_N - p_N^* - (p_T - p_T^*)$ is the relative price of nontraded to traded goods between the countries. In contrast to the standard Balassa-Samuelson model, the 'relative-relative price' of nontraded to traded goods between countries is not equally a function of the deviations in relative productivities, as it is a function of relative differences in sectoral markups. These two drivers, however, obviously have different influences on the equilibrium real exchange rate in a more complete model, because productivity directly increases output as well as relative prices, while the wage markups do not.

The importance of the *relative* difference of price markups is intuitively clear. If the Home country has 10% higher markups than the Foreign country in both sectors, prices will be higher by 10%, ceteris paribus. But the relative price of nontraded goods, a key driver of the real exchange rate, will not be different, since prices of *both* traded and nontraded goods are higher by the same proportion.

We may then ask whether this implies that labor market imperfections have no influence on the real exchange rate in the case when $\mu_T^* - \mu_T - (\mu_N^* - \mu_N) = 0$, that is, when there are *no sectoral* but only *national* differences in firm markups. It turns out that such direct effect also exists, irrespective of whether sectoral wage markups differ, but it is observationally equivalent to the effects of the relative disutility of labor $\chi - \chi^*$. Algebraically, this can be seen from a combination of first order conditions. In logarithms, we can write the implicit labor supply condition as $w^R - q = \sigma c^R + \psi n^R + \chi^R$ where '$.^R$' denotes a value of a Home relative to Foreign variable, expressed in the same currency when necessary. Applying the complete risk sharing condition, this reduces to $w^R = \psi n^R + \chi^R$. We can then use the firm's first order conditions (in either sector) to substitute for w^R, yielding (after substituting for p_N^R):

$$\frac{1}{1-\gamma\kappa} q + a_N^R - \mu^R = \psi n^R + \chi^R$$

where we assume $\mu_N^R = \mu_T^R = \mu^R$. This condition is the only place in the model where μ^R as well as χ^R enter. Consequently, if we define $\tilde{\chi}^R \equiv \chi^R - \mu^R$ we can solve the log-linearized model in the same manner as without labor markups by writing $\tilde{\chi}^R$ instead of χ^R. Then, by construction, the coefficient on μ^R in model's solution (for any variable) must equal the negative of that variable's coefficient on $\tilde{\chi}^R$.

As already reported in Section 3, the general form of the model (assuming no home bias) can be solved for real exchange rate as follows:

$$q = \alpha_\chi \chi^R + \alpha_T a_T^R + \alpha_N a_N^R + \alpha_{\mu_N} \mu_N^R + \alpha_{\mu_N - \mu_T} (\mu_N^R - \mu_T^R)$$

where

$$\begin{aligned}
\alpha_\chi = \alpha_{\mu_N} &= \frac{\sigma(1-\gamma\kappa)}{B} \\
\alpha_{a_T} &= \frac{\sigma(1-\gamma\kappa)}{B} \gamma\kappa\psi(\kappa\lambda + \phi(1-\kappa) - 1) \\
\alpha_{a_N} &= -\frac{\sigma(1-\gamma\kappa)}{B} [1 + \psi(1 + \gamma\kappa(\kappa\lambda + \phi(1-\kappa) - 1))] \\
\alpha_{\mu_N - \mu_T} &= \frac{\sigma(1-\gamma\kappa)}{B} \gamma\kappa\psi(\kappa\lambda + \phi(1-\kappa))
\end{aligned}$$

and

$$B = \sigma + \psi \left(1 + \kappa \left[\sigma(\psi - \theta) + \gamma^2 \kappa(1 - 2\sigma\theta) + \gamma(\sigma(\phi + 2\theta + \kappa(\lambda - \phi - \psi + \theta)) - 2)\right]\right)$$

Under a standard calibration[30] yields coefficients: $\alpha_{\chi} = \alpha_{\mu_N} = 0.22,\ \alpha_{a_T} = 0.26,\ \alpha_{a_N} = -0.71,\ \alpha_{(\mu_N - \mu_T)} = 0.33$.

[30] Specifically, when $\sigma = 2$, $\kappa = 0.6$ (so that the distribution sector accounts for 40% of retail tradable goods in equilibrium), $\theta = 0.7$, $\gamma = \omega = 0.5$, $\Psi = 1$, $\phi = 0.25$ and $\lambda = 8$. See Berka et al. (2018).