Determinants of Relative Sectoral Prices: The Role of Demographic Change

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Abstract

Demographic change raises demand for non-tradable old-age related services relative to tradable commodities. This demand shift increases the relative price of non-tradables and thereby causes real exchange rates to appreciate. We claim that the change in demand affects prices via imperfect intersectoral factor mobility. Using a sample of 15 OECD countries, we estimate a robust increase of relative prices. According to our main estimate, up to one fifth of the average increase in relative prices between 1970 and 2009 can be attributed to population ageing. Further findings confirm the relevance of imperfect factor mobility: Countries with more rigid labour markets experience stronger price effects.

JEL Classification: F14, F31, J11

Keywords: Demographic change, Relative price of non-tradables, Real exchange rate

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1 Introduction

The relative price of non-tradable services to tradable commodities is well-known to be an important determinant of real exchange rates. According to the famous Balassa-Samuelson hypothesis, which dates back to 1964, movements in these relative sectoral prices can be attributed to sectoral differences in productivity growth. Empirical studies tend to find support in favour of the hypothesis.[1] Further determinants of the relative price beyond the Balassa-Samuelson effect operate over the demand-side of the economy. The literature discusses non-homothetic preferences (Bergstrand 1991), government demand (De Gregorio et al. 1994, Galstyan and Lane 2009) and net foreign assets (Lane and Milesi-Ferretti 2002, 2004, and Christopoulos et al. 2012).[2]

In this paper, we propose a country's demographic structure as an additional economic fundamental for the relative price of non-traded goods and we study this relationship empirically. Figure 1 highlights the importance of this determinant. As a point of reference, the left panel depicts the cross-sectional relation between changes in relative prices of non-tradables and productivity growth differentials between tradables and non-tradables for a set of industrialized countries. The strong positive correlation illustrates the Balassa-Samuelson effect. The right panel plots relative-price changes against the average growth rates of old-age dependency ratios (hereafter named OADR), which are defined as the fraction of population aged 65+ to the population of age 15-64. This highlights our proposed channel: changes in the age structure of the population are positively correlated with the growth rate of the relative price of non-tradables. In particular, countries with stronger growth of the OADR experience higher growth in the relative price.

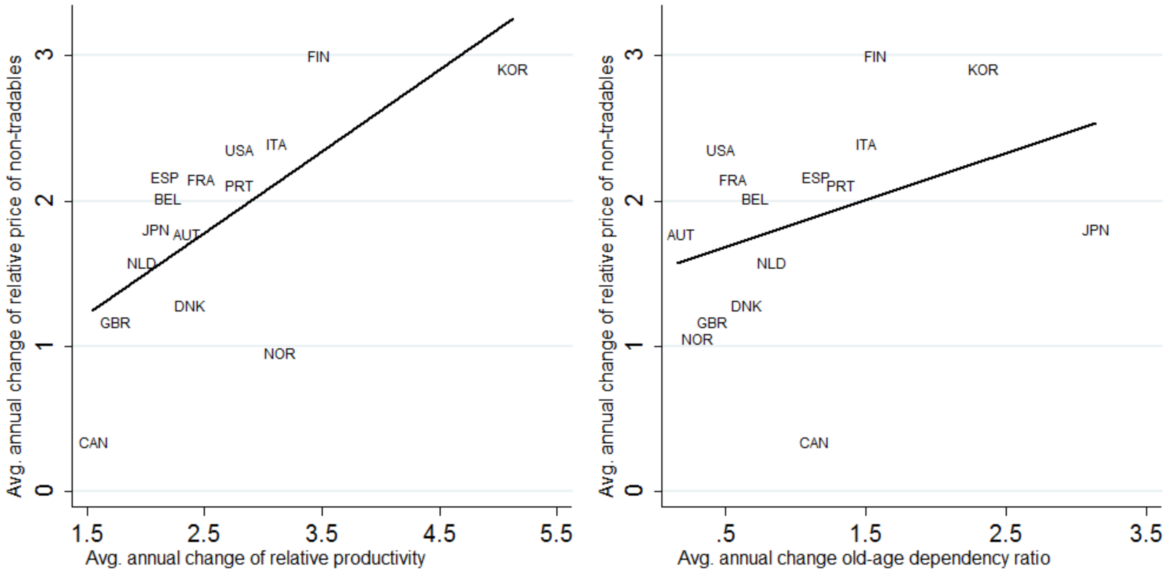
There are several mechanisms how ageing can lead to higher relative prices. In this paper, we focus on the following demand effects. We present evidence that elderly people consume more non-traded services relative to people in working age. This implies an increase in overall demand for those goods due to population ageing. At the same time, the old-age population has lower saving rates than younger cohorts, such that aggregate savings of an ageing society decline (see for instance Higgins 1998 and Yoon et al. 2014), while aggregate consumption increases. Likewise this rise in spending is also biased towards non-tradable goods. If the additional demand for non-traded services of an ageing society is not fully met by higher supply, the relative price of non-tradables increases. We claim that persistent imperfect intersectoral mobility of production factors hampers a reallocation of factor inputs to the non-tradable sector. Since we are concentrating on

[1] See, amongst others, Canzoneri et al. (1999), Kakkar (2003), Berka et al. (2014), and Coto-Martinez and Reboredo (2014), where the latter also consider the role of imperfect competition.

[2] An interesting synopsizing study is conducted by Ricci et al. (2013).

Figure 1: Cross-Sectional Correlations of Relative Price Changes

(a) Relative Productivity (b) Old-age Dependency Ratio



Notes: Average annual changes for 15 OECD countries between 1970 and 2009. Abscissa left panel: productivity in tradable relative to non-tradable sector. Abscissa right panel: old-age dependency ratio (population aged 65+ divided by population aged 15-64). Details on the construction of all variables are given in Appendix 7. Country codes are explained in Table 1.

OECD countries with highly developed capital markets and since the production of non-traded services tends to be labour-intensive, labour market rigidities are – as we show – most important. The empirical literature supports this reasoning. For instance Wacziarg and Wallack (2004), Lee and Wolpin (2006), and Artuc et al. (2010) all provide evidence of substantial and long-lasting intersectoral worker immobility in response to labour demand shocks.[3]

Demographic change may also influence relative prices by other channels, which our analysis accounts for but which we argue to be of less importance. Besides its aforementioned impact on savings, ageing can also influence national investment and, hence, net foreign asset positions. Changes in the latter may imply relative price shifts due to the classical transfer effect. For instance, Lane and Milesi-Ferretti (2002, 2004) argue that higher net foreign assets can generate wealth effects, which lower labour supply. The resulting upward pressure on wages can lead to relative price increases if one sector is relatively labour-intensive, which is usually assumed to be the case for non-tradable services. Simulation results by Krueger and Ludwig (2007) show that changes in net foreign asset positions due to demographic change remain small for the United States and Europe,

[3] Cardi and Restout (2015) demonstrate the importance of labour market rigidities for the transmission of the Balassa-Samuelson effect.

though. A further consequence of ageing could be an increased scarcity of labour relative to capital. This may also result in an upward pressure on wages with corresponding effects on the relative price, given the non-tradable sector is relatively labour-intensive. According to Krueger and Ludwig (2007), Ludwig et al. (2012) and Heijdra and Reijnders (2012) the impact of demographic change on factor prices is expected to be limited.[4] Nevertheless, our empirical analysis considers these channels as well.

To illustrate the relation between sectoral prices and population ageing and in order to provide guidance for the subsequent empirical analysis, we construct a stylized small open economy model with overlapping generations (OLG). We assume two production sectors. As is well known at least since Rogoff (1992), a precondition for any demand effects to matter for relative prices in such a setting is a deviation from full intersectoral factor mobility.[5] Otherwise, supply factors would just shift to the sector that experiences the increase in demand, leaving relative goods prices unchanged. The majority of the related literature merely assumes fixed amounts of sectoral production, thereby ruling out any kind of factor mobility. For instance, Rose et al. (2009) in a related model rely on the strong assumption of completely inelastic labour supply, both in the aggregate and between sectors.[6] In contrast, our model features an endogenous labour supply decision of households, and we explicitly allow for different degrees of imperfect intersectoral mobility of labour. We obtain two testable key results from our model. First, an increase in the old-age dependency ratio leads to an increase in the relative price of non-tradables. The reason is that workers do not reallocate their labour as much as needed to let supply keep up with changing demand. Second, we show that price effects are more pronounced for higher degrees of labour market rigidity.

The basic econometric specification arises from the theoretical model and shows that relative prices depend on the old-age dependency ratio, which is the regressor of main interest. To analyse whether imperfect labour mobility is relevant for the transmission of the effect, we introduce interactions of indices of labour market rigidity with the OADR. We construct a panel of 15 OECD countries that are followed from 1970 to 2009. De-

[4] The literature shows that endogenous human capital adjustments, increased labour market participation of women, migration, international capital flow adjustments, as well as pay-as-you-go pension systems financed by payroll taxes all dampen the effect of demographic change on factor prices. In fact, the number of employees has risen in most countries over the past decades.

[5] Instead of assuming imperfect factor mobility, introducing deviations from the assumption of a small open economy or diminishing returns to scale in at least one sector (see Galstyan and Lane (2009) for an example) is also sufficient for both supply and demand factors to matter in determining relative prices and real exchange rates.

[6] Cantor and Driskill (1999) analyse the effect of a change in death rates on the real exchange rate in a stylized OLG model and find that the direction of the effect depends on the long-run net foreign asset position of a country. However, they do not consider non-tradable goods and their model also does not feature endogenous production.

tailed sector-specific data is classified into tradables and non-tradables to construct relative prices and productivities. To quantify labour market immobility we use the index by Botero et al. (2004) that includes measures of institutional flexibility of the labour market. Our estimation strategy explicitly takes into account the non-stationarity and cross-sectional dependence of the data. To this end, we use the method by Pesaran (2006) and Kapetanios et al. (2011), which finds increasing use in macro panel studies.

Our results indicate a significant link between population ageing and relative sectoral prices. A one percent increase of the old-age dependency ratio inflates the relative price of non-tradables by 0.34 percent. This implies that about one fifth of the average increase in relative prices between 1970 and 2009 can be attributed to the increase of the OADR. The results indicate that the aforementioned demand effects of ageing are predominantly responsible for this finding. Moreover, we identify labour market rigidity as the driving force for the transmission of this demand effect. While price effects are close to zero for countries with very flexible labour markets, they increase monotonically with the degree of rigidity. Various robustness checks underpin the validity of our findings, and demonstrate the importance of this transmission mechanism relative to other possible channels. Further results widen the analysis to the whole population age structure.

We add to the literature that studies the determination of relative sectoral prices and real exchange rates by proposing a demand effect induced by population ageing. Few studies have analysed demographic change in this context. Complementary to our study, Rose et al. (2009) examine the effect of fertility on the real exchange rate arguing that a higher share of the dependent young population leads to lower savings and a higher demand for non-tradables. They confirm their theoretical prediction by finding a depreciating effect of declines in fertility on real exchange rates. Bettendorf and Dewachter (2007) analyse the impact of changes in the whole population age structure on the relative price of non-tradables, but their empirical findings remain in parts insignificant and inconclusive. Using a structural model, Aloy and Gente (2009) focus on Japan and find that declines in population growth are able to explain a large share of the real appreciation since 1970. Andersson and Österholm (2005, 2006) perform reduced-form regressions of real exchange rates on the population age structure. They show that using demographic data can improve forecasts of real exchange rates. For a detailed overview on this literature, see Hassan et al. (2011). Recently, Yoon et al. (2014) and Juselius and Takáts (2015) study the effect of demographic change on inflation using post-war data with mixed results.

Overall, the above mentioned literature studying demand effects on relative prices and real exchange rates is silent about the precise mechanism through which changes in demand translate into price effects.[7]

Our contribution to the literature is, hence, twofold. First, by making use of recent advances in statistical methods, we are able to establish population ageing as a demand-driven determinant of relative sectoral prices. Second, in our empirical specification we pay particular attention to the theoretical transmission mechanism of the old-age related demand effects by introducing labour market rigidity. We show that labour market rigidities are indeed a driving factor for the transmission.

The rest of the paper proceeds as follows. In Section 2, we present stylized facts about age-specific consumption patterns of tradable and non-tradable goods. Section 3 lays out the theoretical model and derives two testable implications. Section 4 translates the theory into an econometric model, describes the data and explores data properties. Section 5 presents the main results of the paper, while various sensitivity checks are shown in Section 6. We conclude the paper in Section 7. Additional results are provided in a supplementary online appendix.

2 Ageing and Consumption

The demand effects of demographic change on sectoral prices rely on the premise that the elderly consume a higher fraction of non-tradable services than the population in working age does. Micro studies on the United States and some European countries all detect this age pattern in consumption data. Hobijn and Lagakos (2003), Börsch-Supan (2003) as well as van Ewijk and Volkerink (2012) present cross-sectional overviews of consumption-age profiles of several different expenditure groups for the U.S., Germany and the Netherlands, respectively. Lührmann (2005, 2008) investigates consumption-age profiles by means of panel data from Germany and the U.K. that enable her to control for all kinds of cohort-, time-, income-, and household-effects. The essence of these studies is that when people become older, they tend to reduce their expenses on tradable goods categories like 'transportation', 'furniture and home electronics' and 'clothing', while demand for non-tradables, such as 'housing' and 'health care goods and services' increases. Based on their findings, Hobijn and Lagakos yet discuss the introduction of an additional CPI for the elderly in the U.S. that takes into account their differing consumption spending patterns.

[7] Although the importance of certain transmission channels like non-homothetic preferences, decreasing returns to scale or production factor immobility is generally acknowledged theoretically, it is usually not further investigated in the empirical specifications, cf. De Gregorio et al. (1994), Galstyan and Lane (2009), and Lane and Milesi-Ferretti (2002, 2004) and Rose et al. (2009), for example. One exception are Christopoulos et al. (2012), who explicitly evaluate the importance of frictions on capital markets in developing countries for real exchange rate determination both theoretically and empirically.

However, household data does not cover the full scope of changes in consumption patterns at the aggregate level. In particular, it does not take into account the substantial public spending on health and long-term care. According to OECD data, average health care spending of member states amounts to about ten percent of GDP in recent years, of which on average only 30 percent are financed by the private sector. Hagist and Kotlikoff (2005) estimate age profiles of health care spending for a sample of ten OECD countries and show that expenditures at old age are a multiple of those in working age. For instance, average health care expenses already double between the age groups 50-64 and 65-69.

As an illustration we quantify the difference between aggregate tradable and non-tradable consumption shares at working age and during retirement exemplarily for a specific country. We choose the United States in the year 2011. To this end, we combine micro data of the U.S. Consumer Expenditure Survey (CE) with aggregate data on Medicare and Medicaid health care spending. First, we map the numerous expenditure categories in the CE data on the tradable and non-tradable sector, based on sector classifications by De Gregorio et al. (1994), to obtain expenditures per capita on tradable and non-tradable goods for young (aged 15-64) and old people (aged 65+).[8] Next, we add Medicare and Medicaid spending per capita, which are classified to be non-tradable, to consumption expenditures of the two age groups. The resulting share of aggregate non-tradable consumption of the older people (83 percentage points) is on average about eight percentage points higher than in case of the younger people (75 percentage points). The magnitude of the difference depends on the chosen country and time period, though. Braude (2000) and van Ewijk and Volkerink (2012) conduct similar exercises for the United States in 1990 and the Netherlands in 2010 respectively and quantify the difference in shares to be, even larger, about 20 (70 versus 50) and 13 (70 versus 57) percentage points. In sum, the available evidence suggests substantial differences in aggregate consumption shares of tradables and non-tradables over the life cycle. Changes in the age distribution of the population are therefore expected to induce non-negligible changes in aggregate demand.

3 Model

We employ a model with overlapping generations and two production sectors $j = \{T, N\}$ to study the effect of population ageing on the relative price of non-tradable goods. Following the literature on structural real exchange rate determination, we assume a small open economy, where interest rates are taken as given by world markets.

A continuum of households lives for at most two periods, in each period t a young

[8] Regarding details on data sources, the reader may be referred to Appendix 7.

and an old generation is alive. Every young individual faces a probability π_t of growing old. The population size of the young generation is normalized to unity. Therefore, π_t is the ratio of old relative to young households. It can be interpreted both in an aggregate perspective as the OADR and in an individual perspective as the life expectancy of the household in period t for $t+1$. In order to capture the observations made in the last section in an abstract fashion, young households receive utility from the consumption of tradable commodities C_t^T and disutility of labour effort L_t, whereas the elderly enjoy the consumption of non-tradable services C_t^N.[9] Households maximize lifetime utility given by

$$U\left(C_t^T, L_t\right) + \beta\pi_t U\left(C_{t+1}^N\right), \tag{1}$$

where $\beta \in (0,1)$ is the subjective discount factor. Utility in working age is given by $U\left(C_t^T, L_t\right) = \ln C_t^T - \ln L_t$, while utility of the elderly is given by $U\left(C_{t+1}^N\right) = \ln C_{t+1}^N$. The assumed preference structure is chosen to obtain analytical solutions and implies both an intertemporal elasticity of substitution and a Frisch-elasticity of one. Labour supply can be allotted to both sectors of production. Following Horvath (2000) and Cardi and Restout (2015), households have a preference to work in both sectors, which drives a wedge between sectoral wages. Total labour in the utility function is defined by the CES-aggregate

$$L_t = \left[\left(L_t^T\right)^{\frac{\rho+1}{\rho}} + \left(L_t^N\right)^{\frac{\rho+1}{\rho}}\right]^{\frac{\rho}{\rho+1}}, \tag{2}$$

where L_t^j denotes hours worked in the tradable ($j = T$) and non-tradable ($j = N$) sector respectively. $\rho > 0$ measures the elasticity of substitution between labour supplies in both sectors. For $\rho \to \infty$, hours worked are perfect substitutes and workers would devote all working time to the sector that pays the highest wage, while for $\rho < \infty$, workers have a preference to diversify their labour supply and are willing to work in both sectors even in the presence of wage differentials. Hence, ρ measures the degree of imperfect intersectoral labour mobility, where small values of ρ imply less mobility. In line with empirical evidence cited above, this modelling choice generates persistent labour market frictions and shall be regarded as a short-cut for more comprehensive models of labour market rigidities, in order to allow for explicit analytical solutions and comparative statics.

Temporal budget constraints are given by

$$C_t^T = L_t^T W_t^T + L_t^N W_t^N - S_t \tag{3}$$

[9] A generalized setting in which both generations consume both types of goods does not change the results qualitatively as long as preferences of the elderly are biased in favour of non-tradables.

and

$$P_{t+1}C_{t+1}^{N} = \frac{(1+R^{*})}{E_{t}\pi_{t}}S_{t}. \tag{4}$$

W_t^T and W_t^N label wages in the two sectors, S_t denotes household savings that are invested on international capital markets and that yield an exogenously given return of R^*, which is the world interest rate. In addition, we assume a perfect annuity market, where assets of those who deceased are passed to the survivors, so that the return on savings is $(1+R^*)/\pi_t$.[10] Finally, we let P_t denote the relative price of non-tradable services to tradable commodities and assume the price of tradable goods to be given by world markets and normalized to unity.[11] The first-order conditions of the household's maximization problem yields the standard Euler-equation, $C_{t+1}^N/C_t^T = \beta(1+R^*)/P_{t+1}$, and an equation on how to optimally supply labour in the two sectors, given by

$$\frac{L_t^T}{L_t^N} = \left(\frac{W_t^T}{W_t^N}\right)^{\rho}. \tag{5}$$

Condition (5) states that relative hours worked depend on the wage ratio $\left(W_t^T/W_t^N\right)$ and the elasticity of substitution ρ.

Both in the tradable and the non-tradable sector a homogeneous consumption good is produced by perfectly-competitive firms using labour L_t^j and physical capital K_t^j as inputs within the Cobb-Douglas technology

$$Y_t^j = F\left(A_t^j, K_t^j, L_t^j\right) = A_t^j\left(K_t^j\right)^{\alpha_j}\left(L_t^j\right)^{1-\alpha_j}, \tag{6}$$

where Y_t^j and A_t^j are output and productivity in sector $j = \{T, N\}$, respectively. Firms borrow capital on international markets, which is assumed to fully depreciate within one period. In this small open economy setting, profit maximization and perfect competition among firms yield that optimal sectoral capital intensities are tied down by productivity and the world interest rate, while real sectoral wages depend on exogenous parameters and exogenous stochastic processes for A_t^j and π_t only.

A competitive equilibrium in this economy is defined as a sequence of prices and quantities such that optimality conditions of all agents hold and markets clear in each period for a given interest rate R^*, a given price of tradable goods and sectoral productivities A_t^j. All agents operate as price takers. Households choose consumption C_t^T and C_{t+1}^N, savings

[10] Alternatively, one could enrich the model by a warm-glow bequest motive. This would imply less dissaving in old age, but at the same time already more savings in young age due to anticipation effects. In sum, effects on the consumption path of households are limited or even absent.

[11] As discussed repeatedly in the literature (for instance in Froot and Rogoff 1995), the relative price of non-tradables is directly related to the real exchange rate in such a setting.

S_t and sectoral labour supplies L_t^j, while firms decide on their labour and capital demand $\left(L_t^j, K_t^j\right)$ in both sectors. Labour markets clear every period. Savings of households are fully invested at international capital markets and firms borrow all capital from abroad. Supply of tradable commodities Y_t^T has to equal domestic demand C_t^T plus net exports. The market clearing condition of the non-tradable sector is given by

$$Y_t^N = \pi_t C_t^N. \tag{7}$$

In steady state, the relative price of non-tradables can be shown to evolve according to

$$P = k \left(\frac{\left(A^T\right)^{\frac{1-\alpha_N}{1-\alpha_T}}}{A^N} \right) \left(\frac{\pi \varkappa}{1 - \pi \varkappa} \right)^{\frac{1-\alpha_N}{1+\rho}}, \tag{8}$$

where k and $\varkappa$ are positive constants and $P > 0$ as long as $1 - \pi\varkappa > 0$.[12] Accordingly, P is driven by two main components in this framework. The first term of (8) in parentheses illustrates the well-known Balassa-Samuelson effect. It states that an increase in productivity in the tradable sector generates an increase in the relative price of non-tradables and appreciation of the real exchange rate, while productivity growth in the non-traded sector yields a decline in the relative price and real depreciation. The second component of the equation, on which this paper lays its focus, highlights the effect of population ageing as well as the role of labour market rigidities in its transmission on the relative price. In particular, we are able to show the following:

Proposition 1 *In steady state, the effect of ageing on the relative price of non-tradables is positive:*

$$\frac{\partial P}{\partial \pi} > 0 \tag{9}$$

Proposition 2 *In steady state, the effect of ageing on relative prices is the higher, the lower intersectoral labour mobility,*

$$\frac{\partial \left(\partial P / \partial \pi\right)}{\partial \rho} < 0, \tag{10}$$

Increasing the old-age dependency ratio leads to higher demand for non-tradable services: Currently young households will consume less tradables and save more for old-age. Once old, they consume the proceedings of their higher savings in form of more non-tradable goods. In case of perfect factor mobility ($\rho \to \infty$) these demand shifts are fully

[12] Formal proofs of the results in this section are given in the supplementary appendix.

met by higher supply as labour moves immediately to the service sector due to positive wage pressure. Production of services increase until wages are equal in both sectors again, leaving relative goods prices unchanged. The last term of (8) converges to one. In contrast, in case of imperfect intersectoral labour mobility ($\rho < \infty$), higher demand in the non-tradable service sector entails a positive effect on the relative price of non-tradables, since reallocation of labour is not exhaustive: increased demand is only partly met by changes in supply and partly by an increase in the relative price.

4 Estimation Procedure and Data

4.1 Econometric Model

Propositions 1 and 2 constitute the two main hypotheses we intend to test. To this end, we set up a reduced form econometric specification based on (8), which is given by

$$\ln(p_{it}) = c_i + \gamma_1 \ln(oadr_{it}) + \gamma_2 \ln(oadr_{it}) \cdot lri_i + \boldsymbol{\gamma}' X_{it} + u_{it}. \tag{11}$$

Sub-indices denote country i and time period t respectively, c_i labels country-specific intercepts and u_{it} is an error term, whose structure will be discussed below. To allow for a convenient interpretation as (semi-)elasticities, variables are, when sensible, used as natural logarithms. The dependent variable $\ln(p_{it})$ is the natural log of a measure of the relative price of non-tradables. The covariate of main interest, $\ln(oadr_{it})$, is the empirical counterpart of π_t and denotes the log of the old-age dependency ratio. According to our first hypothesis, based on Proposition 1, its coefficient γ_1 should possess a positive sign. Our second hypothesis, deduced from Proposition 2, claims that imperfect labour mobility leads to higher price effects of ageing. This is tested for by including an interaction term of $\ln(oadr_{it})$ and a measure of labour market rigidities, lri_i, which is considered to be the empirical counterpart of ρ. For expositional reasons, we begin with using a binary variable, $\widetilde{lri}_i = \{0, 1\}$, for lri_i in the interaction term. This binary index takes on a value of one for countries with an LRI-value above the sample mean ($\widetilde{lri}_i = \mathbf{1}_{lri_i \geq \overline{lri}_i}$) and a value of zero else ($\widetilde{lri}_i = \mathbf{0}_{lri_i < \overline{lri}_i}$). This way the coefficient γ_1 can be interpreted as the effect of ageing for countries with an LRI-value below mean, while $\gamma_1 + \gamma_2$ indicates the effect for countries with an LRI-value above mean. In a next step, we estimate the interaction effect by using the actual country-specific values of lri_i. In this case, $\partial \ln(p_{it}) / \partial \ln(oadr_{it}) = \gamma_1 + \gamma_2 lri_i$ gives the partial effect of ageing for the respective value of lri_i. According to Proposition 2, the partial effect of the old-age dependency ratio on the relative price is thus expected to be positive and to increase for higher degrees of labour market rigidity.

Further explanatory variables are summarized in the vector X_{it}. Its elements are motivated by (8) and by the related empirical literature. We consider variables, for which there exists broad consensus on their importance in determining real exchange rates and relative prices of non-tradables.[13] First, we include productivity in the tradables relative to the non-tradables sector (relative sectoral productivity) to account for the classic Balassa-Samuelson effect. Next, we add GDP per capita to control for demand-side effects, for instance due to non-homothetic preferences that regard non-tradable services as luxuries and tradable commodities as necessities – an approach proposed first by Bergstrand (1991). Moreover, GDP per capita is capable of capturing effects of changes in factor endowments in the spirit of Bhagwati (1984) as GDP is strongly correlated with the capital-labour ratio of the economy. Higher capital-labour ratios lead to higher wages, in particular in the labour-intensive non-tradable sector and, thereby, to a higher relative price. To this extent, GDP per capita also controls for supply-side effects of demographic change, such as changes in the relative scarcity of labour. In the presence of these effects, the coefficient of GDP is expected to be positive. The third factor we consider is government consumption relative to GDP to control for further demand effects, since public expenditures are known to be biased towards non-tradables. Given this reading, its coefficient should also be positive. Evidence on this effect is for instance provided by Ricci et al. (2013) and Galstyan and Lane (2009). Lane and Milesi-Ferretti (2002, 2004) deal with wealth effects of net foreign asset positions on real exchange rates. According to their argument, which is related to Keynes' classical transfer problem, an increase in net foreign assets induces wealth effects that reduce labour supply. This hits labour-intensive non-tradable sectors relatively stronger, thereby leading to an increase of the relative price of non-tradables. To control for this kind of effect, we augment our specifications with a variable on net foreign assets relative to GDP. As demographic change can influence net foreign asset positions via its effect on savings, this variable can also capture these indirect price effects of population ageing. Since net foreign assets can also attain negative values, it is the only element of X_{it}, which is not used in logs.

By the inclusion of net foreign assets and GDP per capita, we control for the transfer and supply-side effects of demographic change that are discussed in the introduction. As a consequence, the coefficient of the old-age dependency ratio will predominantly capture the demand effects of ageing on relative prices.

[13] Ricci et al. (2013) provide an overview about which variables may belong to this canonical set of determinants.

Table 1: Sample Overview

Country	Abbrev.	Coverage	Country	Abbrev.	Coverage
Austria	AUT	1976-2009	Korea	KOR	1971-2009
Belgium	BEL	1975-2009	Netherlands	NLD	1977-2009
Canada	CAN	1970-2006	Norway	NOR	1970-2009
Denmark	DNK	1970-2009	Portugal	PRT	1977-2006
Finland	FIN	1970-2009	Spain	ESP	1980-2009
France	FRA	1970-2008	United Kingdom	GBR	1971-2007
Italy	ITA	1970-2009	United States	USA	1977-2009
Japan	JPN	1970-2008	*Full Sample (avg.)*	–	1973-2008

Notes: N=546 usable observations in the benchmark model.

4.2 Data Description

The empirical analysis is based upon a new-constructed panel data set of 15 OECD countries with annual observations beginning earliest in 1970 and ending at the latest in 2009. No country is followed for less than 30 years. On average we have 36 annual observations per country. Overall, we command 546 usable observations in the benchmark model. The choice of countries is restricted by the availability of sufficiently detailed data on sectoral prices and productivity over sufficiently long time horizons. An overview of the sample dimensions is given in Table 1. All data stems from publicly available sources, such as the OECD STAN data base or the Penn World Tables. Details on the sources and regarding the construction of all variables are shifted to a data appendix at the end of the paper. A list of all variables used throughout the text and summary statistics can be found in Table 2.

The relative price of non-tradable goods is constructed as the ratio of price indices of the non-tradables and the tradables sector. Vice versa, the relative sectoral productivity refers to the productivity ratio of the tradables relative to the non-tradables sector. As in Canzoneri et al. (1999) and Ricci et al. (2013), sectoral productivities are measured as value added per worker. The old-age dependency ratio, which is one of the standard measures in population economics, is defined as population aged older than 65 divided by population in working age (15-64). The the old-age population share measures the amount of the population aged 65+ relative to total population, while the working-age population share accounts for the ages 15-64 to total population. Similar to the OADR, the young-age dependency ratio (hereafter named YADR) measures the amount of dependent children aged 0-14 relative to the population in working age (15-64). The total fertility rate is defined as the number of children that would be born to a woman during her childbearing years if she bears children in accordance with current age-specific fertility rates.

Table 2: Summary Statistics

	Unit	1970 Mean	1970 Std. Dev.	2009 Mean	2009 Std. Dev.	Average Annual Change[a]
Relative Price on Non-Tradables	Index	63.80	17.28	121.86	18.82	1.86
Old-Age Dependency Ratio	%	15.04	5.25	23.51	4.30	1.06
Old-Age Population Share	% of Pop.	9.57	3.37	15.82	2.63	1.24
Working-Age Population Share	% of Pop.	63.14	4.16	67.29	2.01	0.17
Young-Age Dependency Ratio	%	43.95	1.41	25.12	0.30	-1.37
Total Fertility Rate	%	2.54	0.83	1.65	0.27	-0.84
Labour Market Rigidity Index, LRI	[0,1]	0.53	0.22	0.58	0.17	–
Labour Market Rigidity Index, LRI^{EPI}	[0,6]	2.33[b]	1.19[b]	2.14	0.72	-0.36
Capital Openness, KAOPEN	[0,1]	0.39	0.28	0.96	0.13	3.70
Economic Freedom, EconFree	[0,100]	69.18[c]	6.21[c]	73.26	5.36	0.53
Labour Market Freedom, LabFree	[0,100]	67.83[d]	17.97[d]	65.64	20.06	-0.13
Relative Sectoral Productivity	Index	52.86	21.73	148.75	42.01	2.65
GDP per capita	2005 Int$	14,242	4943	34,650	7425	2.31
Government Consumption	% of GDP	7.36	1.35	7.20	1.18	-0.01
Net Foreign Assets	% of GDP	-1.94	34.85	-1.39	48.94	0.12

Notes: [a]Cross-sectional mean of average annual growth rates in percent. [b]Mean and Std. Dev. in 1985 instead of 1970 due to data limitations. [c]Mean and Std.Dev. in 1995. [d]Mean and Std.Dev. in 2005. The relative price of non-tradables is defined relative to the price of tradables, whereas relative sectoral productivity is defined as productivity in the tradables relative to the non-tradables sector.

Table 2 illustrates the magnitude and evolution of the data over time. As most variables in our sample feature clear upward trends, we present means and standard deviations at the beginning and end of the observation period instead of the less meaningful overall sample statistics.

The labour market rigidity index, LRI, in the main results is taken from Botero et al. (2004). This measure is widely used both in academia and by institutions such as the World Bank. It is defined as the average of four other indices, namely alternative employment contracts, cost of increasing hours worked, cost of firing workers, and dismissal procedures. This composite index can attain values between zero and one, where higher values represent larger rigidities. Table 3 reveals a wide variation of the index in our sample. As one would expect, the index takes on substantially lower values for Anglo-American than for continental European countries (e.g. United States 0.22 versus France 0.74). Yet, a drawback of this measure is that it does not reflect changes of these rigidities over time, since it is merely a fixed number per country. This issue is addressed by the OECD Indicators of Employment Protection. These contain a time-*variant* measure of the strictness of employment protection for years following 1985. We denote this index

Table 3: Labour Market Rigidity (LRI) per Country

Country	LRI	Mean(LRI^{EPI})	Country	LRI	Mean(LRI^{EPI})
Austria	0.50	2.64	Korea	0.45	2.64
Belgium	0.51	1.78	Netherlands	0.73	2.93
Canada	0.26	0.92	Norway	0.69	2.33
Denmark	0.57	2.15	Portugal	0.81	4.68
Finland	0.74	2.43	Spain	0.74	2.83
France	0.74	2.39	United Kingdom	0.28	1.01
Italy	0.65	2.76	United States	0.22	0.26
Japan	0.16	1.67	*Full Sample (avg.)*	0.54	2.23

Notes: LRI denotes Labour Market Rigidity Index and has a range of [0,1], see Botero et al. (2004). LRI^{EPI} has a range of [0,6], see OECD Indicators of Employment Protection. For both indices lower values mean lower degrees of rigidity.

as LRI^{EPI} and use it for a robustness check. This alternative index is on a scale from 0 (least rigid) to 6 (most rigid). As Table 3 reveals, means per country of the LRI^{EPI} yield a similar ranking as the index by Botero et al. (2004). Further indices that measure other economic and legal characteristics (see KAOPEN, EconFree, LabFree in Table 2) will be described within the sensitivity analysis.

Table 4: Cross-Sectional Dependence Tests

	CD_P	$avg.\,(r_{ij})$	$avg.\,(\lvert r_{ij}\rvert)$
Relative Price on Non-Tradables	51.03***	0.86	0.86
Old-Age Dependency Ratio	37.90***	0.64	0.70
Old-Age Population Share	45.73***	0.77	0.78
Working-Age Population Share	27.34***	0.46	0.56
Young-Age Dependency Ratio	51.87***	0.87	0.87
Total Fertility Rate	14.05***	0.23	0.60
Relative Sectoral Productivity	57.92***	0.97	0.97
GDP per capita	58.17***	0.98	0.98
Government Consumption (% of GDP)	10.34***	0.17	0.52
Net Foreign Assets (% of GDP)	-2.48**	-0.041	0.58

Notes: All variables except net foreign assets are measured in logs. CD_P denotes Pesaran (2004) cross-sectional dependence test statistic. Asterisks indicate rejection of the null hypothesis of cross-sectional independence at 10%(*), 5% (**) and 1% (***). $avg.\,(r_{ij})$ and $avg.\,(\lvert r_{ij}\rvert)$ denote average and average absolute cross-sectional correlation coefficients.

4.3 Non-Stationarity and Cross-Sectional Dependence

In order to determine the appropriate estimation techniques, we test the data for cross-sectional correlation and its trend behaviour. Macroeconomic variables are notoriously affected by these two issues, which can seriously distort inference and consistency of estimations.[14]

To check for cross-sectional dependencies, Table 4 presents average (absolute) cross-sectional correlation coefficients and results of Pesaran's (2004) CD_P test statistic, which is $N(0,1)$-distributed under the null hypothesis of cross-sectional independence. The CD_P statistics rejects cross-sectional independence for all variables and the computed average correlation coefficients reveal strong cross-sectional correlations for all variables, except for government consumption and net foreign assets. Altogether, the results leave no doubt that cross-sectional correlation is indeed a problem in this data set.

Table 5: Panel Unit Root Tests

	CIPS	IPS	MW
Relative Price on Non-Tradables	1.27	1.46	24.35
Old-Age Dependency Ratio	5.17	1.94	35.67
Old-Age Population Share	1.11	1.19	34.88
Working-Age Population Share	-4.32***	-2.35***	61.96***
Young-Age Dependency Ratio	-6.68***	-7.70***	49.05**
Total Fertility Rate	1.66	-2.34***	81.66***
Relative Sectoral Productivity	0.46	2.85	25.20
GDP per capita	-0.37	-0.28	22.81
Government Consumption (% of GDP)	1.19	0.032	19.07
Net Foreign Assets (% of GDP)	6.24	5.19	11.30

Notes: All variables except net foreign assets are measured in logs. Results of CIPS (Pesaran 2007), IPS (Im et al. 2003), and MW (Maddala and Wu 1999) panel unit root test statistics. Asterisks indicate rejection of the null hypothesis of a unit root at 10%(*), 5% (**) and 1% (***). Optimal lag length determined by Akaike and Bayesian information criteria searching between 0 and 4 lags.

The trend behaviour of the data is explored by means of three different panel unit root tests. We apply the tests by Im et al. (2003) (IPS), Maddala and Wu (1999) (MW), and by Pesaran (2007) (CIPS). The latter is a panel unit root test of the second generation that can account for cross-sectional correlations, which is important as tests that neglect this issue can have non-negligible size distortions. Results are shown in Table 5. Under the null hypothesis of all three tests, the variable in question entails a unit root, while under the alternative hypothesis at least one series of the panel is stationary. All three tests are based on standard augmented Dickey-Fuller regressions on the individual time series, but in case of the CIPS these are extended by cross-sectional averages in lagged levels and first-differences of the variable in question to address issues of cross-sectional correlation. In

[14] As shown, for instance, by O'Connell (1998) in the context of tests for purchasing power parity, disregarding cross-sectional dependence can come at high costs and may even revert outcomes of empirical studies.

order to control for serial correlation, all tests can include various autoregressive lags. The optimal lag length for each variable is determined by the Akaike and Bayesian information criteria searching between 0 and 4 lags. As Table 5 reveals all variables but the working-age population share, the YADR, (and fertility) are found to be non-stationary. It is, therefore, important to examine the time series properties of the regression residuals as well in order to rule out spurious regression results.

4.4 Econometric Methodology

Given the presence of cross-sectional correlation in the data, we follow Pesaran (2006) in assuming an error term of multi-factorial structure for our panel regression model (11). We describe the error term, u_{it}, by

$$u_{it} = \boldsymbol{\delta}_i' \mathbf{f}_t + \varepsilon_{it}, \quad \varepsilon_{it} \sim \text{i.i.d. } N\left(0, \sigma^2\right), \tag{12}$$

where $\mathbf{f}_t$ is a vector of unobserved, potentially non-stationary common factors, which represent events that appear to influence all countries at the same time. By the vector of individual-specific factor loadings $\boldsymbol{\delta}_i'$, different countries are still allowed to react differently to these common effects. The covariates $x_{it} \in (oadr_{it}, X_{it})$ in (11) can be correlated with the same unobserved factors $\mathbf{f}_t$ as u_{it}, and may be described as a process of the type

$$x_{it} = a_i + \boldsymbol{\eta}_i' \mathbf{f}_t + v_{it}, \tag{13}$$

which is assumed to depend on a fixed effect a_i, the factors $\mathbf{f}_t$ with country-specific factor loadings $\boldsymbol{\eta}_i'$ and a random component v_{it}. In case of economic macro variables, examples for the factors $\mathbf{f}_t$ are common business cycles, the world financial crisis, or the effects of globalization. In case of demographic variables, one may think of changes in working environments, habits, or medical innovations that increase longevity or reduce birth rates such as the contraceptive pill. If common factors are present in u_{it}, but u_{it} and x_{it} are uncorrelated ($\boldsymbol{\delta}_i \neq \mathbf{0}, \boldsymbol{\eta}_i = \mathbf{0}$), error terms in (11) will be cross-sectionally correlated and the use of conventional estimators will yield inefficient standard errors. If u_{it} and x_{it} are correlated additionally ($\boldsymbol{\delta}_i \neq \mathbf{0}, \boldsymbol{\eta}_i \neq \mathbf{0}$), coefficient estimates itself are biased due to a type of omitted variable bias.

A promising approach to remedy the problem is to apply the Common Correlated Effects Pooled (referred to as CCEP) estimator developed by Pesaran (2006). This estimator is practically computed as an ordinary least squares regression, augmented with a set of additional regressors that consists of the cross-sectional averages of the dependent and independent variables, which are interacted with country dummies. As the data are found to be non-stationary, it is likely that at least some components of $\mathbf{f}_t$ are integrated of order one. Kapetanios et al. (2011) show that the CCEP estimator is consistent in presence of unit roots in the unobservable factors. Using Monte Carlo studies, they demonstrate the superiority of the CCEP estimator over other commonly used ones, even in small samples as ours. A further appealing feature of the estimator is that by controlling for the potentially non-stationary common factors the approach helps to deal with the problems of non-stationary data, such as biased inference that can lead to spurious regressions. Given these issues, estimations of (11) are conducted by means of Pesaran's CCEP approach and, for comparison reasons, using the DOLS procedure (see Stock and Watson 1993, Kao and Chiang 2000), which is a widely used methodology for non-stationary panels. It is calculated as a two-way fixed effects model that employs additional leads and lags of first differences of the independent variables.

In order to rule out spurious regressions, we control for stationarity of the regression residuals, again taking into account the possibility of cross-sectional dependence. To this end, we again apply the CIPS test.[15] As is well known, we cannot directly use the critical values from Pesaran (2007) that were constructed for the case of raw data, since regression residuals are calculated as to minimize the sum of their squares. Instead, we generate critical values directly from our sample by applying the Continuous-Path Block Bootstrap method developed by Paparoditis and Politis (2000, 2003), assuming a fixed block length of 10 percent of the overall observation period.[16] This method is explicitly designed to preserve non-stationarity and cross-sectional dependence of the data. We generate 500 bootstrap redraws of the estimated regression residuals. For each redraw we compute the CIPS statistic and thereby generate a distribution of the test statistic.

5 Results

In this section, we provide evidence for the two hypotheses deduced from Propositions 1 and 2. Table 6 presents results of model (11) without (columns I, II) and with (columns III, IV) the interaction term using the CCEP and the DOLS approach.

[15] For instance, Holly et al. (2010) use it for the same purpose.

[16] We adopt and modify code by Fachin (2007), such that it can be used for an unbalanced panel as ours.

Table 6: Main Regressions

Dependent Variable: Relative Price of Non-Tradables	(I) CCEP	(II) DOLS	(III) CCEP	(IV) DOLS
Old-Age Dependency Ratio (OADR)	0.34***	0.20***	0.15	0.20***
	(0.093)	(0.040)	(0.111)	(0.040)
OADR × Labour Market Rigidity (LRI)			0.54***	0.16**
			(0.175)	(0.065)
Relative Sectoral Productivity	0.59***	0.58***	0.58***	0.55***
	(0.046)	(0.034)	(0.045)	(0.038)
GDP per capita	0.41***	-0.054	0.32***	-0.018
	(0.072)	(0.040)	(0.076)	(0.041)
Government Consumption (% of GDP)	0.11**	-0.055	0.073	-0.10**
	(0.051)	(0.039)	(0.051)	(0.043)
Net Foreign Assets (% of GDP)	-0.00083***	-0.0022***	-0.00081***	-0.0020***
	(0.00017)	(0.00017)	(0.00017)	(0.00020)
Residual diagnostics				
CD_P	-2.08**	-3.26***	-2.51**	-3.33***
CIPS	-7.095***	-0.20	-7.21***	0.014
F(OADR,Interaction= 0)			11.91***	16.19***
Observations	546	501	546	501

Notes: All variables except net foreign assets are measured in logs. Regressions based on (11). In regressions (III) & (IV) , OADR is interacted with a binary variable with value one in case of labour market rigidity above average ($\widetilde{lri}_i = \mathbf{1}_{lri_i \geq \overline{lri}_i}$) and zero else ($\widetilde{lri}_i = \mathbf{0}_{lri_i < \overline{lri}_i}$). Country dummies included in all regressions, time dummies in (II) and (IV). Standard errors in parentheses. Asterisks mark significance at 10% (*), 5% (**), 1% (***). Residual diagnostics: CD_P cross-sectional dependence test statistic by Pesaran (2004), Residual stationarity tested by Pesaran's (2007) CIPS test using bootstrapped critical values. F(OADR,Interaction= 0) denotes F-test about joint significance of OADR and the interaction term.

5.1 Ageing and the Relative Price of Non-Tradables

The coefficient of the old-age dependency ratio using the CCEP estimator, cf. column (I), implies that an increase in the OADR by one percent leads to an increase in the relative price of non-tradables of 0.34 percent. The estimate is statistically significant at the 1 percent level and constitutes good evidence for our first hypothesis – population ageing leads to a higher relative price of non-tradables. Since our regression model controls for the transfer and supply-side effects of demographic change, we infer that this finding is mainly driven by changes in demand. Our results complement the empirical finding by

Rose et al. (2009), who find that increases in fertility lead to appreciated real exchange rates.

All of the control variables enter the regression significantly, most of them with both qualitatively and quantitatively reasonable coefficients. An increase in relative sectoral productivity by one percent leads to an increase in the relative price by 0.59 percent, thereby offering evidence in favour of the Balassa-Samuelson effect. The effect of GDP per capita is positive also, which indicates the presence of price effects due to non-homothetic preferences (Bergstrand 1991) or relative factor endowments (Bhagwati 1984), as discussed in Section 4.1. Because of the latter, the variable also captures supply-side effects of demographic change, such as changes in the size of the labour force relative to the capital stock. In line with the related literature, rises in government spending inflate the relative price, though its coefficient is relatively small. As opposed to the intuition given in Section 4.1, the effect of an increase in net foreign assets tends to reduce relative prices, though at a very low rate. An increase in net foreign assets over GDP by one percentage point lowers the relative price by 0.083 percent only. However, according to the findings of Christopoulos et al. (2012), transfer effects are generally negligible for developed countries and only gain importance in case of capital-constraint developing countries, which are not included in our sample. This result also implies that the effects of demographic change on relative prices that operate over the transfer channel are very small. This finding is also in line with simulation results by Krueger and Ludwig (2007).

As diagnostic tests of the regressions, we provide cross-sectional dependence and stationarity tests of the residuals. Cross-sectional independence is rejected at the 5 percent level in column (I), which implies that it is important to consider this type of correlations when testing for residual stationarity. According to the bootstrapped critical values of the CIPS test, a unit root in the residuals can be rejected at the 1 percent level, so that we can rely on the inference of the estimated effects.

In column (II), the CCEP results are contrasted with the same regression using a DOLS estimator.[17] Our finding regarding the effect of ageing on the relative price remains qualitatively unchanged. An increase of the OADR by one percent raises the relative price of non-tradables by 0.20 percent, which is slightly smaller than the CCEP estimate. Turning to the control variables, the effects of relative productivity and net foreign assets can be found close to the CCEP estimates, while the effects of GDP per capita and government consumption turn insignificant. However, in reference to the residual diagnostics, regression (II) appears to be spurious as we are not able to reject the null hypothesis of non-stationarity in the residuals. Though this does not need to imply in-

[17] Due to the limited sample length, we use one lead and lag.

consistent estimates of the coefficients (see Phillips and Moon 1999), inference can be highly misleading and should not be relied on. This difference between the two estimation approaches underlines vividly the capabilities of the CCEP estimator to filter out unobservable non-stationary components of the data.

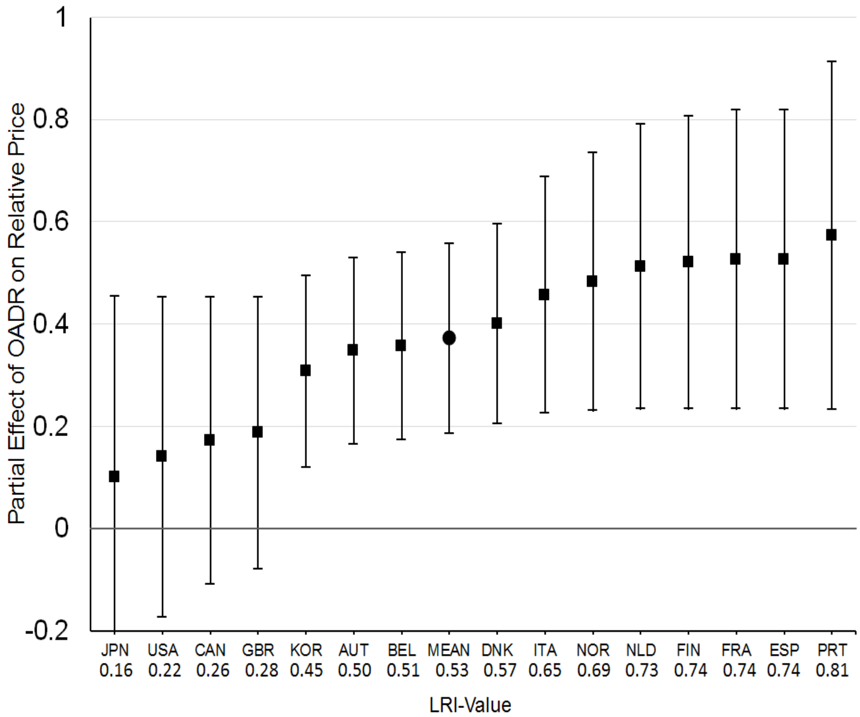
5.2 The Importance of Labour Market Rigidities

Columns (III) and (IV) in Table 6 show results for the regressions augmented with an interaction term of the OADR and the binary variable $\widetilde{lri}_i$ that indicates rigid labour markets. It is constructed with the index by Botero et al. (2004), which is a standard measure in the labour economics literature. In line with our prediction in Proposition 2, the effect of ageing indeed depends on labour market rigidities. In column (III), the coefficient of the OADR now gives the effect of ageing on relative prices for the less rigid countries, where $lri_i < \overline{lri}_i$. For these countries a one percent increase of the OADR implies an increase of the relative price by 0.15 percent, which is less than half the size of the effect in column (I). Statistically this effect is not significantly different from zero. Such findings regularly occur in models with interaction effects due to the inherent multicollinearity between the interaction and its basis variable. The coefficient of the interaction itself is positive and highly significant as theory predicts. The effect of a one percent increase of the OADR for the rigid countries reads 0.15+0.54=0.69 percent. An F-test implies joint significance of the OADR-coefficient and the interaction term at the 1 percent level. In sum, the estimates imply that ageing has a small insignificant effect in less rigid countries, while the effect is strong and positive for rigid countries. In the DOLS model in column (IV) the results are qualitatively similar, though the coefficients are again slightly different from the CCEP estimation. For the countries with less rigid labour markets, the effect of ageing reads 0.20 percent and for those with rigid labour markets it rises to 0.36 percent. In terms of control variables and residual diagnostics both regressions behave similarly to their counterparts in (I) and (II). Again, inference in the DOLS regression has to be questioned due to a rejection of residual stationarity.

To further investigate the effect of the interaction, we reestimate model (11) using the untransformed country-specific values of the labour market rigidity index and the CCEP estimator. By means of this model, we calculate the partial effect of ageing $\partial \ln(p_{it}) / \partial \ln(oadr_{it}) = \gamma_1 + \gamma_2 lri_i$ for the LRI-values of each country in the sample. The resulting effects of the old-age dependency ratio for each country along with their 95 percent confidence band are plotted in Figure 2.[18] Confirming the theoretical predictions

[18] In order to compute significance levels of the partial effect $\partial \ln(p_{it}) / \partial \ln(oadr_{it})$, the standard approach is to reparameterize model (11) by subtracting the respective value of the rigidity index for

Figure 2: Visualisation of Interaction Effects



Notes: Partial effects of OADR, $\partial \ln(p_{it}) / \partial \ln(oadr_{it}) = \gamma_1 + \gamma_2 lri_i$, and corresponding 95% confidence band evaluated at the country-specific LRI-value using model (11). For the underlying regression, see Table 1 in Appendix B. Country codes are explained in Table 1 in the text, LRI values for each country depicted on the abscissa.

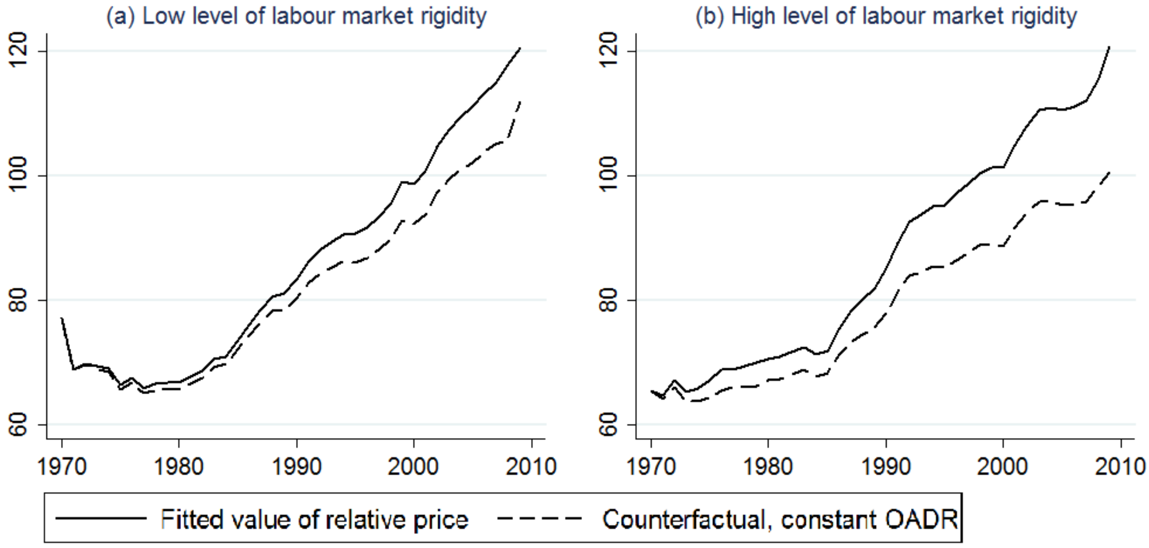
of Proposition 2, the countries with the most flexible labour markets, appearing on the left side of the figure, undergo small price effects of ageing, while countries with higher degrees of rigidity experience larger effects. In particular, for LRI-values up to 0.3 that are related with Anglo-American countries, price effects are estimated to be not statistically different from zero. In case of LRI-values about 0.7 - 0.8, which correspond to (Southern) European countries as France, Spain and Portugal, relative price effects rise up to nearly 0.60 percent. These findings underpin the empirical relevance of Proposition 2 and support the validity of the proposed transmission via imperfect labour market mobility.

5.3 The Economic Significance of Ageing

In terms of economic significance, the estimate of our main regression in column (I) of Table 6 implies the following. As the average old-age dependency ratio in our sample

country j before using it in the interaction term. Model (11) is then given by $\ln(p_{it}) = c_i + \widetilde{\gamma}_1 \ln(oadr_{it}) + \gamma_2 \ln(oadr_{it}) \cdot (lri_i - lri_j) + \boldsymbol{\gamma}' X_{it} + u_{it}$. The coefficient of OADR can then be shown to be the partial effect of OADR for country j, i.e. $\widetilde{\gamma}_1 = \gamma_1 + \gamma_2 lri_j$. This way standard errors can be directly read out to evaluate statistical significance of the partial effect at the respective rigidity value of country j.

Figure 3: Fitted Values and Counterfactual Experiments



Notes: Fitted and counterfactual values of the relative price of non-tradables, based on model (11), for two groups of countries. Left panel (a): Austria, Belgium, Canada, Japan, Korea, United Kingdom, United States. Right panel (b): Denmark, Finland, France, Italy, Netherlands, Norway, Portugal, Spain.

increased from a value of 15 to 23.5 (see Table 2) by about 56 percent between 1970 and 2009, the coefficient of the OADR of 0.34 percent implies a surge of relative prices due to demographic change of about 19 percent. As in the same time period the relative price increased on average by 91 percent, about one fifth of the price movements can be attributed to population ageing. Hence, the effect of ageing appears to be of reasonable and non-negligible magnitude.

Economic significance is further studied by means of a set of counterfactual experiments, presented in Figure 3. The figure provides fitted values (solid lines) of the relative price of non-tradables, as implied by model (11), and compares these to a counterfactual scenario, where the OADR is kept constant at its 1970 value (dashed lines). The difference between the solid and the dashed line can be interpreted as the ceteris paribus effect of ageing on the relative price over the sample period. This is done for two groups of countries. In Panel (a) of Figure 3 the average LRI-value of countries whose LRI is below sample mean are used for calculating fitted values and the counterfactual simulation, while Panel (b) employs the average of countries with LRI values above mean. The figure confirms the relative importance of population ageing and the transmission via labour market frictions. While in both groups relative price effects would be dampened in case of a constant age structure, the difference between actual and counterfactual lines is by far more pronounced for countries with a high degree of labour market rigidity. In numbers, in Panel (a) the difference in the increase of the solid (57 percent) and the dashed line (45 percent) sums up to about 12 percentage points between 1970 and 2009, while in Panel (b) the difference is about 31 percentage points. Simulations for the whole sample and all individual countries can be found in Section B.1 of the supplementary appendix and also support the relevance of ageing for relative prices.

6 Sensitivity Analysis

To provide further evidence of the importance of demographic change for sectoral prices and the proposed channel via labour market rigidities, we analyze various alternative specifications. First, we evaluate the effect of changes in the whole demographic structure on relative prices instead of using the aggregated old-age dependency ratio. Second, we provide robustness checks for our proposed transmission mechanism. To this end, we use an alternative labour market rigidity index, as well as indices that account for further economic and legal factors. The result of the latter exercise is that other potential transmission mechanisms do not have the same empirical support as labour market rigidities do.

6.1 Disentangled Age Effects

To complement the finding of an appreciating effect of young cohorts on the real exchange rate by Rose et al. (2009), we test for the effect of the young-age dependency ratio (the fraction of population aged 0-14 to the population of age 15-64) on the relative price of non-tradables. The intuition for a higher share of the population at young-age is analogue to the impact of the OADR: the young dependent population consumes a higher share of non-tradable goods – such as education – with an upward effect on relative prices. Column (I) of Table 7 shows the results using the CCEP estimator. Consistent with the intuition provided by Rose et al. (2009) for real exchange rates, the young-age dependency ratio is also found to increase the relative price of non-tradables. A 1 percent increase of YADR inflates the relative price by 0.19 percent.[19]

The old-age dependency ratio as the explanatory variable of main interest in Section 5 subsumes two opposing effects. Being defined as the ratio of the retired to the working-age population, this variable simultaneously captures the effect of changes in two population shares. According to our hypothesis both age groups should yield opposite effects on

[19] We also successively added two different demographic measures of young cohorts, the YADR and the fertility rate, to our main model in column (I) of Table 6. The effect of OADR on relative prices remains qualitatively the same in all cases. The coefficient of the YADR also stays close to the one estimated in Table 7. The fertility rate, however, is not found to have a significant effect on relative prices. Full results on these checks can be found in Section B.2 of the supplementary appendix.

relative sectoral prices: A higher share of the working-age population is expected to generate a negative impact due to a relatively lower demand for non-tradables.[20] On the contrary, a higher share of the elderly relative to total population should have a positive impact. To study the impact of a countries' demographic structure on the relative sectoral price in more detail, we decompose the old-age dependency ratio in its components, and use the share of population in working age (15-64) and the old-age population share (65+) relative to total population instead of the OADR.

Table 7: Alternative Demographic Variables

Dependent Variable: Relative Price of Non-Tradables	(I)	(II)	(III)
Young-Age Dependency Ratio	0.19**		
	(0.10)		
Working-Age Population Share		-1.62***	
		(0.34)	
Old-Age Population Share			0.26***
			(0.096)
Relative Sectoral Productivity	0.56***	0.50***	0.63***
	(0.046)	(0.045)	(0.046)
GDP per capita	0.40***	0.51***	0.30***
	(0.067)	(0.071)	(0.067)
Government Consumption (% of GDP)	0.13***	0.16***	0.099*
	(0.049)	(0.048)	(0.052)
Net Foreign Assets (% of GDP)	-0.0005***	-0.0005***	-0.0008***
	(0.00017)	(0.00017)	(0.00018)
Residual diagnostics			
CD_P	-2.39**	-3.00***	-1.19
CIPS	-8.958***	-8.48***	-7.43***
Observations	546	546	546

Notes: All variables except net foreign assets are measured in logs. Method of estimation: CCEP. Regressions based on (11) without interaction term using alternative demographic variables. Country dummies included in all regressions. Standard errors in parentheses. Asterisks mark significance at 10% (*), 5% (**), 1% (***). Residual diagnostics: CD_P cross-sectional dependence test statistic by Pesaran (2004), Residual stationarity tested by Pesaran's (2007) CIPS test using bootstrapped critical values.

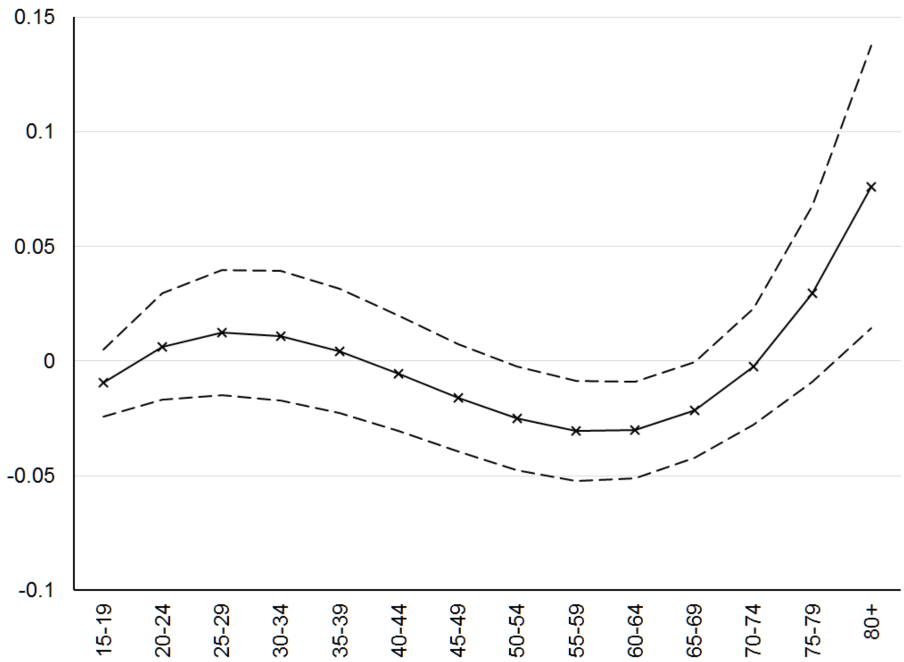
Columns (II) and (III) of Table 7 present results of using the two alternative demographic variables. In line with theory, the working-age population share has a negative

[20] In the model of Section 3 this would imply a decline in π_t.

and significant coefficient in column (II). A 1 percent increase leads to a decline of relative prices by 1.62 percent. Column (II) shows – similar to the results in Table 6 – that population ageing significantly inflates relative prices. An increase of the old-age population share by 1 percent increases relative prices by 0.26 percent. The quantitative importance of the relatively high coefficient of the working-age population share is put into perspective by the fact that changes in the working-age share are far less pronounced than in the old-age population share (see Table 2).

The coefficients of the control variables in all three columns are close to those estimated in our main regression in Section 5. Inference of the regressions is valid, since non-stationarity of the residuals can be rejected. The CD_P test indicates presence of cross-sectional correlations in (I) and (II), but not in (III). Regression results using the DOLS estimator are again close to CCEP estimates, but residual stationarity has to be rejected following bootstrapped critical values to the CIPS test. Therefore, all results using the DOLS methodology are relegated to Section B.2 of the supplementary appendix.

Figure 4: Approximated Coefficients for 5-year Age-Bins



Notes: Coefficients ν_l for 5-year age bins approximated with the cubic age polynomial according to equations (14) and (15). Dashed lines represent the 95% confidence band. The inverted S-shape implies negative coefficients for the working-age population and positive coefficients for older ages.

Having established the opposing effects of the different age groups, we now investigate the effect of the demographic structure in more detail. To this end, we group the population in smaller bins. In particular, we construct $\mathcal{L} = 14$ population groups of 5-year intervals starting at age 15 until age 80+, where the last age bin covers all households at ages 80 and older. For each country i and time period t we compute the fraction of the age interval relative to the total population at ages 15 onwards. The age variables enter the estimation equation as $\sum_{l=1}^{\mathcal{L}} \nu_l \cdot age_{lit}$, where age_{lit} is the population share of age bin l in country i at period t and ν_l is the corresponding coefficient. Because of our relatively small sample we approximate this detailed demographic information by an age polynomial as it is done by Fair and Dominguez (1991) and Higgins (1998). In particular, we assume that the coefficients ν_l lie along a cubic polynomial[21]

$$\nu_l = \omega_0 + \omega_1 l + \omega_2 l^2 + \omega_3 l^3. \tag{14}$$

We can estimate ω_1, ω_2 and ω_3 by introducing auxiliary age variables Z_{it} in an estimation model related to (11) as follows:[22]

$$\ln(p_{it}) = c_i + \sum\nolimits_{k=1}^{3} \omega_k Z_{kit} + \boldsymbol{\gamma}' X_{it} + u_{it}, \tag{15}$$

where the Z_{kit} are defined as

$$Z_{kit} = \sum_{l=1}^{\mathcal{L}} l^k \cdot age_{lit} - \frac{1}{\mathcal{L}} \sum_{l=1}^{\mathcal{L}} l^k \sum_{l=1}^{\mathcal{L}} age_{lit} \text{ with } k = 1, 2, 3;$$

and X_{it} is again a vector of additional control variables. Once the coefficients ω_1, ω_2 and ω_3 are estimated, we can approximate the coefficients ν_l for each age bin l by equation (14). Since the age bin coefficients ν_l are linear transformations of our estimated coefficients ω_k, we can employ the delta method to compute standard errors and confidence bands for ν_l. The results of our estimation are depicted in Figure 4. A full set of results from regression (15) including coefficients for the auxiliary demographic variables Z_{kit} and residual diagnostics are relegated to Section B.2 of the supplementary appendix.

The approximated coefficients ν_l of the age bins are in line with our theory. Coefficients for age bins during working age are negative from ages 40 onwards. At older ages past 70, the coefficients turn positive. Overall, 5 out of 14 age bins are significant at the 5

[21] We have to approximate the demographic structure by a polynomial in order to mimimize the losses of degrees of fredom. For each additional explanatory variable the CCEP estimator involves interaction terms of cross-sectional averages with all country dummies. Hence, three addional variables that are required for a cubic polynomial already imply 48 additional variables in the estimation.

[22] Imposing a zero-sum restriction for ν_l, $\sum_{l=1}^{\mathcal{L}} \nu_l = 0$, implies $\omega_0 = -\frac{1}{\mathcal{L}} \left[\omega_1 \sum_{l=1}^{\mathcal{L}} l + \omega_2 \sum_{l=1}^{\mathcal{L}} l^2 + \omega_3 \sum_{l=1}^{\mathcal{L}} l^3 \right]$. This removes the perfect colinearity between the constant and the age shares in the regression.

percent level.[23] A test for joint significance of all Z_{kit} reveals that demography is jointly significant at the 1 percent level. To evaluate the economic significance, we employ the same counterfactual experiment as in Section 5 and thus keep the whole age structure constant at their values in 1970. We find that demographic changes can account for 17 percentage points of the increase in the relative price between 1970 and 2009, which is of the same order of magnitude as in our main specification.

6.2 Robustness of the Transmission Mechanism

To further validate our proposed transmission channel we present results of a further set of robustness analyses. First, in our main regression, we constructed the interaction term using the old-age dependency ratio and the labour market rigidity index by Botero et al. (2004) to study the transmission mechanism of demand effects. Although widely used in the literature, a shortcoming of this index is that it is constant over time and therefore cannot reflect changes in these rigidities due to, e.g., labour market reforms. In this section, we employ the time-varying labour market rigidity index by the OECD, denoted by LRI^{EPI}, instead. In a second exercise, we perform a different estimation strategy to test for the influence of labour market rigidities. We split the sample into countries with high and low degrees of rigidity, using the measure of Botero et al. (2004) again, and run separate regressions on the subsamples. We assign countries to the low rigidity sample, whose index value is below the sample mean, $\overline{\text{LRI}}$, and to the high rigidity sample else. A third group of results aims at testing whether other factors besides labour market rigidities, such as capital market frictions or imperfect competition, could also drive or be confounded with the transmission of the age effects.

Results on the first two points are shown in Table 8 where the CCEP estimator is used for all regressions, while results from the DOLS method are relegated to Section B.2 of the supplementary appendix. Column (I) shows results of using the time-varying labour market rigidity index. Due to limited data availability for the LRI^{EPI}, the sample size is reduced by about one third to 360 observations. Both the coefficient of the OADR and of the interaction term enter the regression statistically significant. The coefficients can now be interpreted as follows. The partial effect of the OADR on the relative price is given by $\partial \ln(p_{it}) / \partial \ln(oadr_{it}) = 0.42 + 0.016 \cdot lri_{it}^{EPI}$. For instance at the sample mean of the rigidity index ($\overline{lri}^{EPI} = 2.23$), the partial effect is 0.46 percent, which is of the same order

[23] We also tried different degrees for the polynomials. A quadratic polynomial yields similar results with respect to the shape (u-shaped) and significance of the coefficients. 4th and 5th order polynomials yield a similar shape but mostly insignificant coefficients for the age bins. As explained before, this might be due to the loss of too many degrees of freedom associated with the inclusion of additional variables.

Table 8: ALTERNATIVE LRI AND SAMPLE SPLIT

Dependent Variable: Relative Price of Non-Tradables	(I) Alternative Variable	(II) Sample Split	(III)
		LRI $<$ $\overline{\text{LRI}}$	LRI $\geq$ $\overline{\text{LRI}}$
Old-Age Dependency Ratio	0.42**	0.17*	0.53***
	(0.17)	(0.090)	(0.10)
OADR $\times$ Labour Market Rigidity (LRI^{EPI})	0.016**		
	(0.0079)		
Relative Sectoral Productivity	0.27***	0.36***	0.75***
	(0.081)	(0.077)	(0.058)
GDP per capita	0.45***	0.039	0.19**
	(0.11)	(0.083)	(0.088)
Government Consumption (% of GDP)	0.22***	0.086	-0.0067
	(0.065)	(0.061)	(0.075)
Net Foreign Assets (% of GDP)	-0.000086	-0.0012***	-0.00093***
	(0.00020)	(0.00039)	(0.00019)
Residual diagnostics			
CD_P	-1.52	-2.95***	1.11
CIPS	-6.26**	-4.49***	-5.46***
F(OADR,Interaction= 0)	4.45**		
Observations (countries)	360	254 (7)	292 (8)

Notes: All variables except net foreign assets are measured in logs. Method of estimation: CCEP. Regressions based on (11). Country dummies included in all regressions. Interpretation of interaction in column (I): Partial effect of ageing is given by $0.42+0.016 \cdot lri_{it}^{EPI}$. The low rigidity set (column II) contains Austria, Belgium, Canada, Japan, Korea, United Kingdom, and United States, while the high rigidity set (column III) covers Denmark, Finland, France, Italy, Netherlands, Norway, Portugal, and Spain. Standard errors in parentheses. Asterisks mark significance at 10% (*), 5% (**), 1% (***). Residual diagnostics: CD_P cross-sectional dependence test statistic by Pesaran (2004), Residual stationarity tested by Pesaran's (2007) CIPS test using bootstrapped critical values. F(OADR,Interaction= 0) denotes F-test about joint significance of OADR and the interaction term.

of magnitude as the result shown in Figure 2 (0.37 percent). Evaluating the effect of the OADR at different points of LRI^{EPI} also yields qualitatively analogue outcomes to those shown in Figure 2, albeit the differences between countries are smaller.

Columns (II) and (III) show results for the two subsamples with high and low rigidities, respectively. As hypothesized, the effect of ageing is very marked for the rigid countries, LRI$\geq$ $\overline{\text{LRI}}$, and about 50 percent larger than the effect found for the full sample in Table 6. For the subsample with lower degrees of rigidity, LRI$<$ $\overline{\text{LRI}}$, the coefficient of the old-age dependency ratio is about half the size of the coefficient in Table 6. The effects of the controls are in all cases broadly comparable to those found earlier, though government consumption enters the regressions insignificantly when splitting the sample. In terms of residual diagnostics, non-stationarity is again rejected at high rates for all regressions.

Table 9: Tests for other Transmission Mechanisms

Dependent Variable:	(I)	(II)	(III)	(IV)
Relative Price of Non-Tradables	Capital Market Openness	Economic Freedom	Labor-Market Freedom	All Interactions
Old-Age Dependency Ratio	0.31***	0.37***	0.51***	0.54***
	(0.09)	(0.13)	(0.12)	(0.16)
OADR × KAOPEN	-0.0036			-0.0041
	(0.0032)			(0.0032)
OADR × EconFree w/o Labour		-0.044		-0.097
		(0.18)		(0.18)
OADR × LabFree			-0.36**	-0.38**
			(0.17)	(0.18)
Relative Sectoral Productivity	0.61***	0.59***	0.55***	0.57***
	(0.047)	(0.045)	(0.05)	(0.051)
GDP per capita	0.36***	0.40***	0.41***	0.35***
	(0.074)	(0.08)	(0.071)	(0.081)
Government Consumption (% of GDP)	0.091*	0.10**	0.09*	0.074
	(0.051)	(0.052)	(0.051)	(0.052)
Net Foreign Assets (% of GDP)	-0.00078***	-0.00082***	-0.00079***	-0.00075***
	(0.00017)	(0.00017)	(0.00017)	(0.00017)
Residual diagnostics				
CD_P	-2.33**	-2.12**	-2.36**	-2.68***
CIPS	-7.33**	-7.10***	-6.82***	-7.07***
F(OADR,Interactions= 0)	5.79***	6.88***	9.10***	4.10***
Observations	542	546	546	542

Notes: Method of estimation: CCEP. All variables except net foreign assets are measured in logs. Regressions based on (11). In each column OADR is interacted with binary variables with value one in case the respective index value is above average and zero else. Country dummies included in all regressions. Standard errors in parentheses. Asterisks mark significance at 10% (*), 5% (**), 1% (***). Residual diagnostics: CD_P cross-sectional dependence test statistic by Pesaran (2004), Residual stationarity tested by Pesaran's (2007) CIPS test using bootstrapped critical values. F(OADR,Interaction= 0) denotes F-test about joint significance of OADR and the interaction terms.

In Table 9 we evaluate if other factors besides labour market rigidities are also relevant for the size of the price effects of population ageing.

Similar to our proposed mechanism via labour market rigidities, capital market frictions could impair intersectoral and international adjustments of inputs after demand shifts due to population ageing. To test for the relevance of this channel, we re-estimate our main result from Table 6 using the index of capital market openness (KAOPEN) by Chinn and Ito (2006) instead of the LRI. The KAOPEN index measures restrictions on cross-border financial transactions on a scale from 0 to 1, where higher values imply more open capital markets. Results using the index can be found in column (I) of Table 9. As with the LRI, the interaction is constructed with a binary variable that takes on a value of one in case of capital market openness above average and zero else. The coefficient of OADR is highly significant and of the same size as in our main regressions. The effect of ageing does not change for different degrees of capital market rigidities though: the interaction term is insignificant and close to zero.

In presence of imperfect competition, demand shifts could lead to relative price changes, even with completely elastic supply. To test for such a mechanism and to cover a broad set of potential other factors, we employ the Index of Economic Freedom (EconFree) by Miller and Kim (2015) next. This index is based on 4 categories, each consisting of several sub-dimensions. The categories are: rule of law (including items like property rights and corruption), limited government (taxation, government spending), regulatory efficiency (business, market entry, and labour market regulation; monetary policy), and open markets (trade policy, banking and investment regulation). Each sub-dimension is graded on a scale of 0 to 100, where higher values mean more economic freedom. The index itself is an unweighted average of all sub-dimensions. To avoid confounding effects with labour market rigidities, we construct a sub-index with all components of the EconFree excluding labour market regulation. Regression results using a binary interaction term are shown in column (II) of Table 9. The OADR is found to be highly significant and of similar size as in the benchmark results, while the interaction with the Economic Freedom index remains small and insignificant. Countries with less general regulation do not experience significantly different price effects of ageing. In opposition, using an interaction between the labour market regulation component (LabFree) of the EconFree in column (III) confirms our previous finding of a significant effect of the OADR that decreases in case of less labour market regulation (note that the interaction effect is now negative since higher values mean less regulation for the LabFree). Adding all 3 interaction terms – LabFree, EconFree, and KAOPEN – jointly in column (IV) does not alter these findings. In supplementary appendix B.2 we show that our baseline result from column (III) in Table 6, i.e. using the labour market rigidity index from Botero et al. (2004) instead of LabFree, is also not affected by the inclusion of interaction terms with additional rigidity indices.

We value this as further evidence for our proposed transmission mechanism: higher labour market rigidity implies stronger effects of population ageing on the relative price of non-tradables, even after controlling for frictions and regulations on other markets, which on itself do not play any apparent role.

7 Conclusion

This paper analyzes the impact of demographic change on the relative price of non-tradable services to tradable commodities. We illustrate by means of a simple model how changes in demand induced by population ageing can affect relative prices. Imperfect labour market mobility is key for the transmission of changes in demand in this setup. We test these hypotheses empirically for a panel of 15 OECD countries. By making use of the CCEP estimator, we account for non-stationary and cross-sectional dependence, which is present in our data. Our results indicate a statistically and economically significant relation of reasonable size between the old-age dependency ratio and the relative price of non-tradables, which implies an appreciation of the real exchange rate in case of population ageing. Our model predicts that about one fifth of relative price movements between 1970 and 2009 can be attributed to demographic change. We further find support for our proposed transmission mechanism through labour market rigidities. In particular, Southern European countries like Portugal, Spain and France with more rigid labour markets experience stronger price effects due to population ageing than Anglo-American countries that feature lower degrees of rigidity.

This paper extends the literature on structural real exchange rate determination by offering the demographic structure of the population as a further complementary explanation for international inflation differentials apart from existing ones like relative productivities or government spending. Furthermore, we stress the role of labour market imperfections for the transmission of effects on relative prices and real exchange rates. As trends in population ageing for the countries of our sample are forecasted to exacerbate in the upcoming decades, one can expect considerable price changes due to demographic change.

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Appendix

Data Sources and Construction

An overview of all data sources is given in Table 11. In case of all variables but relative prices and productivities no further data transformations (except for taking logarithms) are needed. In case of these two exemptions, the variables have to be constructed by hand from raw data. Below, we describe the procedure to construct the relative price and productivity measures, which is based on De Gregorio et al. (1994). The Structural Analysis (STAN) database by the OECD publishes detailed production data of some of its member states, where total value added is decomposed into nine standardized sectors. Series are provided both in current and constant prices using the base year 2000, allowing the calculation of sectoral deflators. In order to classify sectors to be tradable or non-tradable, De Gregorio et al. compute average ratios of exports to production for each sector. If this measure exceeds a given threshold, they use 10 percent, a sector is marked as tradable. These classifications are still used frequently, for instance by Ricci et al. (2013), and we also stick to it. An overview of all sectors with their original notation by the OECD and their classification of tradability are given in Table 10. Accordingly, five sectors, accountable for 65 percent of total value added in the year 2000, are classified as non-tradable, the four remaining sectors as tradable. As one can see, all service sectors except for '*Transport, Storage and Communications*' that is accountable for only 6.7 percent of total value added, are marked as non-tradable.

Table 10: Sector Classifcations

Sector	Share of Value Added	Classification
Agriculture, hunting, forestry and fishing	3.2	T
Mining and quarrying	0.3	T
Manufacturing	24.8	T
Electricity, gas and water supply	3.0	N
Construction	7.0	N
Wholesale and retail trade - restaurants and hotels	15.0	N
Transport, storage and communications	6.7	T
Finance, insurance, real estate and business services	22.9	N
Community, social and personal services	17.1	N

Notes: Share of Value Added in % based on own calculations, defined as unweighted cross-sectional average over whole sample in 2000 using data in constant prices. N and T denote non-tradability and tradability, respectively. Classifications are taken from De Gregorio et al. (1994).

To obtain the relative price of non-tradables, we first compute separate price indices for non-tradable services and tradable commodities using the following formula:

$$p^j = \frac{\sum_{s\in j} VALU^s}{\sum_{s\in j} VALK^s} \text{ for } j = \{T, N\}\,, \tag{16}$$

where s is an index running over all sub-sectors in sector j, and $VALU$ and $VALK$ denote value added in current and constant prices, respectively. Subsequently, the deflator of non-tradables is divided by its counterpart of tradable goods to obtain the relative price $p_{it} = p_{it}^N / p_{it}^T$, which is – after taking logs – employed in the regressions. Data on relative productivity, which we compute as value added per worker, also stems from the STAN database. First, productivity measures for both the non-tradable and the tradable sector are calculated by dividing sectoral value added at constant prices ($VALK$) by sectoral total employment ($EMPN$):

$$spr^j = \frac{VALK^j}{EMPN^j} \text{ for } j = \{T, N\} \tag{17}$$

Relative sectoral productivity, rpr_{it}, as used in the regression analysis is then constructed by $rpr_{it} = spr_{it}^T / spr_{it}^N$ and taking logs of the result.

Table 11: DATA SOURCES OVERVIEW

Type of Data / Variable	Source	Name in Source
OECD Macro Health Care Data	OECD Health Data: http://www.oecd.org/health/health-systems/ oecdhealthdata.htm	HCTOT-PARPIB-HFTOT, HCTOT-PARTOT-HF2
Consumer Expenditure Survey 2011	http://www.bls.gov/cex/	Various
Medicare spending per capita	Health and Health Care of the Medicare Population 2009: http://www.cms.gov/Research-Statistics-Data-and-Systems/ Research/MCBS/Data-Tables.html	Table 4.1
Medicaid spending per capita	http://www.census.gov/compendia/statab/cats/health_nutrition/ medicare_medicaid.html	Table 151
Relative prices and productivities	OECD - Structural Analysis (STAN): http://stats.oecd.org/	VALU, VALK, EMPN
Demographic data	United Nations - World Population Prospects (WPP): http://esa.un.org/wpp/	Various
Labour market rigidity index (LRI)	Botero et al. (2004): http://faculty.tuck.dartmouth.edu/rafael-laporta/research-publications	index_labor7a
Labour market rigidity index (LRI^{EPI})	OECD Indicators of Employment Protection: www.oecd.org/employment/protection	EPRC_V1
Capital market openness index	Chinn and Ito (2006): http://web.pdx.edu/~ito/Chinn-Ito_website.htm	KAOPEN
Index of Economic Freedom	Miller and Kim (2015): http://www.heritage.org/index/	Various
GDP per capita, government consumption (% of GDP)	Heston et al. (2012) - Penn World Table (PWT) 7.1: http://pwt.econ.upenn.edu/	rgdpl, kg
Net foreign assets (% of GDP)	Lane and Milesi-Feretti (2007): http://www.philiplane.org/EWN.html	NFA/GDP

References

Aloy, M. and Gente, K. (2009). The Role of Demography in the Long-Run Yen/USD Real Exchange Rate Appreciation, *Journal of Macroeconomics* **31**(4): 654–667.

Andersson, A. and Österholm, P. (2005). Forecasting Real Exchange Rate Trends Using Age Structure Data - the Case of Sweden, *Applied Economics Letters* **12**(5): 267–272.

Andersson, A. and Österholm, P. (2006). Population Age Structure and Real Exchange Rates in the OECD, *International Economics Journal* **20**(1): 1–18.

Artuc, E., Chaudhuri, S. and McLaren, J. (2010). Trade Shocks and Labor Adjustment: A Structural Empirical Approach, *American Economic Review* **100**(3): 1008–1045.

Balassa, B. (1964). The Purchasing-Power Parity Doctrine: A Reappraisal, *Journal of Political Economy* **72**(6): 584–596.

Bergstrand, J. H. (1991). Structural Determinants of Real Exchange Rates and National Price Levels: Some Empirical Evidence, *American Economic Review* **81**(1): 325–334.

Berka, M., Devereux, M. B. and Engel, C. (2014). Real Exchange Rates and Sectoral Productivity in the Eurozone, *NBER Working Paper 20510.*

Bettendorf, L. and Dewachter, H. (2007). Ageing and the Relative Price of Nontradables, *Tinbergen Institute Discussion Paper 2007-064/2.*

Bhagwati, J. N. (1984). Why are Services Cheaper in the Poor Countries?, *Economic Journal* **94**(374): 279–286.

Botero, J., Djankov, S., Porta, R., Lopez-De-Silanes, F. C. and Shleifer, A. (2004). The Regulation of Labor, *Quarterly Journal of Economics* **119**(4): 1339–1382.

Braude, J. (2000). Age Structure and the Real Exchange Rate, *Bank of Israel Discussion Paper 2000.10.*

Börsch-Supan, A. (2003). Labor Market Effects of Population Aging, *Labour* **17**(Special Issue): 5–44.

Cantor, R. and Driskill, R. (1999). Effects on the Real Exchange Rate of Demographic and Fiscal Policy-induced Changes in National Savings: Is the Conventional Academic Wisdom Always Right? mimeo.

Canzoneri, M. B., Cumby, R. E. and Diba, B. (1999). Relative Labor Productivity and the Real Exchange Rate in the Long Run: Evidence for a Panel of OECD Countries, *Journal of International Economics* **47**(2): 245–266.

Cardi, O. and Restout, R. (2015). Imperfect Mobility of Labor Across Sectors: A Reappraisal of the Balassa-Samuelson Effect, *Journal of International Economics* **97**(2): 249–265.

Chinn, M. D. and Ito, H. (2006). What Matters for Financial Development? Capital Controls, Institutions, and Interactions, *Journal of Development Economics* **81**(1): 163–192.

Christopoulos, D. K., Gente, K. and León-Ledesma, M. A. (2012). Net Foreign Assets, Productivity and Real Exchange Rates in Constrained Economies, *European Economic Review* **56**(3): 295–316.

Coto-Martinez, J. and Reboredo, J. C. (2014). The Relative Price of Non-Traded Goods under Imperfect Competition, *Oxford Bulletin of Economics and Statistics* **76**(1): 24–40.

De Gregorio, J., Giovannini, A. and Wolf, H. C. (1994). International Evidence on Tradables and Nontradables Inflation, *European Economic Review* **38**(6): 1225–1244.

Fachin, S. (2007). Long-Run Trends in Internal Migrations in Italy: A Study in Panel Cointegration with Dependent Units, *Journal of Applied Econometrics* **22**(2): 401–428.

Fair, R. C. and Dominguez, K. M. (1991). Effects of the Changing U.S. Age Distribution on Macroeconomic Equations, *American Economic Review* **81**(5): 1276–1294.

Froot, K. A. and Rogoff, K. (1995). Perspectives on PPP and Long-Run Real Exchange Rates, *in* G. M. Grossman and K. Rogoff (eds), *Handbook of International Economics*, Vol. 3, Elsevier, chapter 32, pp. 1647–1688.

Galstyan, V. and Lane, P. R. (2009). The Composition of Government Spending and the Real Exchange Rate, *Journal of Money, Credit and Banking* **41**(6): 1233–1249.

Hagist, C. and Kotlikoff, L. (2005). Who's Going Broke? Comparing Growth in Healthcare Costs in Ten OECD Countries, *NBER Working Paper 11833.*

Hassan, A., Salim, R. and Bloch, H. (2011). Population Age Structure, Saving, Capital Flows and the Real Exchange Rate: a Survey of the Literature, *Journal of Economic Surveys* **25**(4): 708–736.

Heijdra, B. J. and Reijnders, L. S. M. (2012). Human Capital Accumulation and the Macroeconomy in an Ageing Society, *CESifo Working Paper 4046*.

Heston, A., Summers, R. and Aten, B. (2012). Penn World Table Version 7.1, Center for International Comparisons of Production, Income and Prices at the University of Pennsylvania, http://pwt.econ.upenn.edu/.

Higgins, M. (1998). Demography, National Savings, and International Capital Flows, *International Economic Review* **39**(2): 343–69.

Hobijn, B. and Lagakos, D. (2003). Social Security and the Consumer Price Index for the Elderly, *FRB of New York: Current Issues in Economics and Finance* **9**(5): 1–6.

Holly, S., Pesaran, M. H. and Yamagata, T. (2010). A Spatio-Temporal Model of House Prices in the USA, *Journal of Econometrics* **158**(1): 160–173.

Horvath, M. (2000). Sectoral Shocks and Aggregate Fluctuation, *Journal of Monetary Economics* **45**(1): 69–106.

Im, K. S., Pesaran, M. H. and Shin, Y. (2003). Testing for Unit Roots in Heterogeneous Panels, *Journal of Econometrics* **115**(1): 53–74.

Juselius, M. and Takáts, E. (2015). Can Demography Affect Inflation and Monetary Policy?, *Bank for International Settlements Working Paper 485*.

Kakkar, V. (2003). The Relative Price of Nontraded Goods and Sectoral Total Factor Productivity: An Empirical Investigation, *Review of Economics and Statistics* **85**(2): 444–452.

Kao, C. and Chiang, M.-H. (2000). On the Estimation and Inference of a Cointegrated Regression in Panel Data, *Advances in Econometrics* **15**: 179–222.

Kapetanios, G., Pesaran, M. H. and Yamagata, T. (2011). Panels with Non-Stationary Multifactor Error Structures, *Journal of Econometrics* **160**(2): 326–348.

Krueger, D. and Ludwig, A. (2007). On the Consequences of Demographic Change for Rates of Returns to Capital, and the Distribution of Wealth and Welfare, *Journal of Monetary Economics* **54**(1): 49–87.

Lane, P. R. and Milesi-Ferretti, G. M. (2002). External Wealth, the Trade Balance, and the Real Exchange Rate, *European Economic Review* **46**(6): 1049–1071.

Lane, P. R. and Milesi-Ferretti, G. M. (2004). The Transfer Problem Revisited: Net Foreign Assets and Real Exchange Rates, *Review of Economics and Statistics* **86**(4): 841–857.

Lane, P. R. and Milesi-Ferretti, G. M. (2007). The External Wealth of Nations Mark II: Revised and Extended Estimates of Foreign Assets and Liabilities, 1970-2004, *Journal of International Economics* **73**(2): 223–250.

Lee, D. and Wolpin, K. I. (2006). Intersectoral Labor Mobility and the Growth of the Service Sector, *Econometrica* **74**(1): 1–46.

Ludwig, A., Schelkle, T. and Vogel, E. (2012). Demographic Change, Human Capital and Welfare, *Review of Economic Dynamics* **15**(1): 94–107.

Lührmann, M. (2005). Population Aging and the Demand for Goods & Services, *MEA Discussion Paper 95-05.*

Lührmann, M. (2008). Effects of Population Ageing on Aggregated UK Consumer Demand. Mimeo.

Maddala, G. S. and Wu, S. (1999). A Comparative Study of Unit Root Tests with Panel Data and a New Simple Test, *Oxford Bulletin of Economics and Statistics* **61**(Special Issue): 631–652.

Miller, T. and Kim, A. B. (2015). Index of Economic Freedom, The Heritage Foundation & The Wall Street Journal.

O'Connell, P. G. (1998). The Overvaluation of Purchasing Power Parity, *Journal of International Economics* **44**(1): 1–19.

Paparoditis, E. and Politis, D. N. (2001). The Continuous-Path Block-Bootstrap, *in* M. L. Puri (ed.), *Asymptotics in Statistics and Probability*, VSP International Science Publishers, pp. 305–320.

Paparoditis, E. and Politis, D. N. (2003). Residual-Based Block Bootstrap for Unit Root Testing, *Econometrica* **71**(3): 813–855.

Pesaran, M. H. (2004). General Diagnostic Tests for Cross Section Dependence in Panels, *IZA Discussion Paper 1240.*

Pesaran, M. H. (2006). Estimation and Inference in Large Heterogeneous Panels with a Multifactor Error Structure, *Econometrica* **74**(4): 967–1012.

Pesaran, M. H. (2007). A Simple Panel Unit Root Test in the Presence of Cross-Section Dependence, *Journal of Applied Econometrics* **22**(2): 265–312.

Phillips, P. C. B. and Moon, H. R. (1999). Linear Regression Limit Theory for Nonstationary Panel Data, *Econometrica* **67**(5): 1057–1111.

Ricci, L. A., Milesi-Ferretti, G. M. and Lee, J. (2013). Real Exchange Rates and Fundamentals: A Cross-Country Perspective, *Journal of Money, Credit and Banking* **45**(5): 845–865.

Rogoff, K. (1992). Traded Goods Consumption Smoothing and the Random Walk Behavior of the Real Exchange Rate, *NBER Working Paper 4119*.

Rose, A. K., Supaat, S. and Braude, J. (2009). Fertility and the Real Exchange Rate, *Canadian Journal of Economics* **42**(2): 496–518.

Samuelson, P. A. (1964). Theoretical Notes on Trade Problems, *Review of Economics and Statistics* **46**(2): 145–154.

Stock, J. H. and Watson, M. W. (1993). A Simple Estimator of Cointegrating Vectors in Higher Order Integrated Systems, *Econometrica* **61**(4): 783–820.

van Ewijk, C. and Volkerink, M. (2012). Will Ageing Lead to a Higher Real Exchange Rate for the Netherlands?, *De Economist* **160**(1): 59–80.

Wacziarg, R. and Wallack, J. S. (2004). Trade Liberalization and Intersectoral Labor Movements, *Journal of International Economics* **64**(2): 411–439.

Yoon, J.-W., Kim, J. and Lee, J. (2014). Impact of Demographic Changes on Inflation and the Macroeconomy, *IMF Working Paper 14/210*.