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# RELATIVE SECTORAL PRICES AND POPULATION AGEING: A COMMON TREND

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# Relative Sectoral Prices and Population Ageing: A Common Trend*

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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 between 1970 and 2009, we estimate a robust increase of relative prices due to population ageing. Further findings confirm the relevance of imperfect factor mobility: Countries with more rigid labour markets experience stronger price effects.

JEL Classification: J11, F41, E39.

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

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# 1 Introduction

The relative price of the non-tradable service sector to the tradable commodities sector is well-known to be an important determinant of real exchange rates. In the absence of changes in the terms of trade, an increase of the relative price of non-tradables directly translates into an appreciation of the real exchange rate. In the long-run, the behaviour of the relative price of non-tradables – and thus of the real exchange rate – can at least in part be explained by economic fundamentals.[1] According to the standard theory, movements in relative sectoral prices are entirely attributed to sectoral differences in productivity growth – the famous Balassa-Samuelson hypothesis, cf. Balassa (1964) and Samuelson (1964). Though this hypothesis tends to be confirmed in applied work[2], empirical evidence suggests as complementary explanations the presence of demand-side effects and differing market structures. Important determinants of the relative price beyond the Balassa-Samuelson effect discussed in the literature are non-homothetic preferences (Bergstrand 1991), government demand (De Gregorio et al. 1994, Galstyan and Lane 2009), net foreign assets (Lane and Milesi-Ferretti 2004, and Christopoulos et al. 2012) and imperfect competition (Coto-Martinez and Reboredo 2012).

In this paper we propose a country's demographic structure as an economic fundamental for explaining the relative price of non-traded goods. Figure 1 highlights the importance of this determinant. The left panel depicts the cross-sectional relation between long-run 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.[3] The right panel plots relative-price changes with 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 also positively correlated with the growth rate of relative prices. In particular, countries with stronger growth of the OADR experience higher growth in the relative price of non-tradable goods.

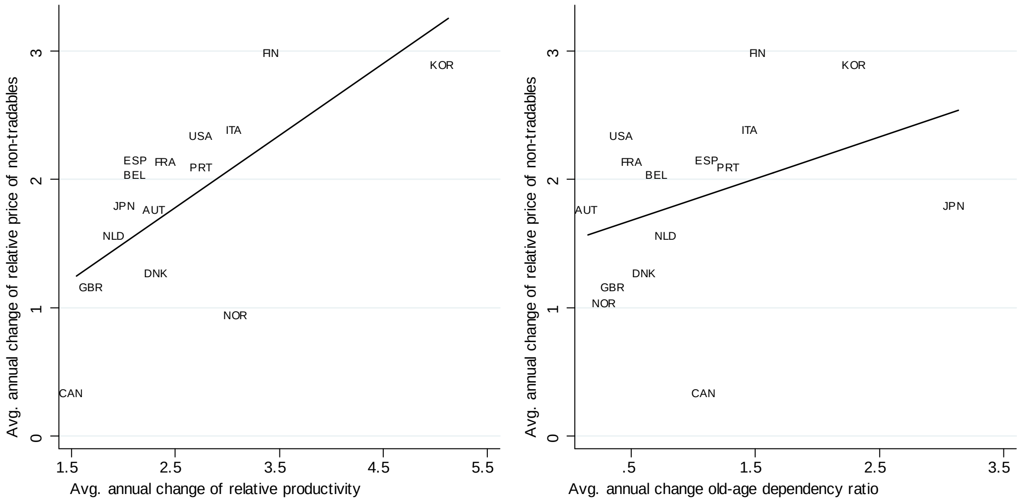
How can population ageing lead to changes in relative sectoral prices? We present evidence that elderly people have a higher demand for non-traded services, such as health care, than people in working age do. The additional demand for services

[1] See Taylor and Taylor (2004) and Froot and Rogoff (1995) for surveys of the literature.

[2] See amongst others Canzoneri et al. (1999) and Kakkar (2003).

[3] This illustration is adapted from the seminal paper of De Gregorio et al. (1994) which has also been used by Obstfeld and Rogoff (1996, Chapter 4), who provide an in-depth discussion of the Balassa-Samuelson model.

Figure 1: Cross-sectional correlations of relative price changes




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 A.1. Country codes are explained in Table 2.

resulting from an ageing society is not fully met by higher supply. 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 developed OECD countries, we presume that labour market rigidities are most important.

In the following, we first construct a stylized overlapping generations (OLG) model to illustrate the channel and to provide guidance for the subsequent empirical analysis. We assume a small open economy with two production sectors. To accommodate frictions on factor markets, we include imperfect intersectoral labour mobility, as proposed by Horvath (2000) and Cardi and Restout (2013). 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 – having a preference for working in both sectors even in the presence of wage differentials – 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 rigidity. The general importance of labour market rigidities is emphasised by Lee and Wolpin (2006), who identify high mobility costs of changing the sector of employment. As well as Cardi and Restout (2013), they provide evidence that suggests substantial long-run labour immobility.

Next, we study the impact of demographic change on relative prices empirically. The basic econometric specification arises from a linearized version of the theoretical model and shows that relative prices depend, aside from sectoral productivity differentials, 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 different indices of labour market rigidity and the OADR. We construct a panel of 15 OECD countries that are followed from 1970 to 2009. The detailed sector-specific data is classified into tradables and non-tradables to construct sectoral 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. As we are interested in long-run relationships our model is estimated in levels while the trend behaviour of all variables as well as the presence of cross-sectional dependence in our data is taken into account.

Our main results indicate a significant link between population ageing and relative sectoral prices. A one percentage point increase of the old-age dependency ratio inflates the relative price of non-tradables by about 0.6 to 2 percent. This implies that up to one fifth of the average increase in relative prices, that almost doubled between 1970 and 2009, can be attributed to the average increase of the OADR. Moreover, we identify imperfect labour mobility as the driving force for the transmission of the demand effect on the relative price: While price effects are close to zero for countries with very flexible labour markets, the effects rise up to four per cent in case of strong rigidities. Various robustness checks underpin the validity of our findings.

The existing literature on the impact of demographic change on sectoral prices and real exchange rates is scarce. Rose et al. (2009) analyse the effect of changes in fertility on real exchange rates in a panel data set of 87 countries between 1975 and 2005. They find that declines in fertility lead to depreciated real exchange rates. Closer to our work, Bettendorf and Dewachter (2007) focus on the effect of changes in the population age structure on the relative price of non-tradables for 16 OECD countries from 1970 to 2002. However, their empirical findings remain inconclusive and insignificant in the majority of cases. Gente (2006) analyses real exchange rates of emerging Asian economies within a calibrated OLG model, in which she relaxes the assumption of perfect international capital market mobility. In this setting real exchange rates do not merely depend on relative productivities, but also on factors like capital borrowing constraints and demographic variables.

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. The theory is translated into an econometric model in Section 4, while Section 5 describes the data and explores data properties. Section 6 presents results of our benchmark specification and of various robustness checks. We conclude the paper in Section 7.

# 2 Ageing and Consumption

The impact of demographic change on sectoral prices relies 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 particularly '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.

However, results from micro data do not cover the full scope of changes in consumption patterns. In particular, they do 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 per cent of GDP in recent years, of which on average only 30 per cent 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.[4] For instance, average health care expenses already double between the age groups 50-64 and 65-69.

[4] The related literature debates to which extent this result is just driven by health care costs incurring in the terminal two years of life (Seshamani and Gray 2004, Zweifel et al. 1999), but even in presence of such an effect, higher numbers of old people are always related with higher aggregate expenditures on health care.

In order to explicitly quantify the difference between tradable and non-tradable consumption shares at working age and during retirement, we combine micro data of the U.S. Consumer Expenditure Survey (CE) from 2011 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+).[5] 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 non-tradable consumption of 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 though depends on the chosen country and time period. 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, differences in consumption shares of tradables and non-tradables over the life cycle are substantial and 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 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, so the interest rate is determined by world markets, i.e. $R_t = R^*$.

Households live for at most two periods, in each period $t$ a young and an old generation is alive. Every young individual faces a probability $\pi_t$ of growing old. The population size of the young generation is normalised to unity. Therefore, $\pi_t$ can be interpreted as the OADR. In order to capture the observations made in the last section in a simple manner, young households receive utility from tradable commodities $C_t^T$ and leisure $(1 - L_t)$, whereas the elderly merely consume non-tradable services $C_{t+1}^N$.[6] Total time being also normalised to unity can be allotted to

[5] Regarding details on data sources, the reader may be referred to Appendix A.1.

[6] A generalised setting in which both generations consume fractions of both types of goods does

working hours in the two sectors and leisure. Households maximize lifetime utility given by

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

where $\beta \in (0,1)$ is the subjective discount factor. Utility in working age is given by $U\left(C_t^T, 1-L_t\right)=\log C_t^T+\phi\log\left(1-L_t\right)$, where $\phi$ is the weighting factor of leisure. Per period utility of the elderly is given by $U\left(C_{t+1}^N\right)=\log C_{t+1}^N$.

Following Horvath (2000) and Cardi and Restout (2013), 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 a 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 $L_t^T$ and $L_t^N$ denote hours worked in the tradable and non-tradable sector respectively. $\rho$ measures the elasticity of substitution between labour supplies in both sectors, i.e. the ease by which labour can be reallocated between the two 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. For $\rho < \infty$, workers have a preference for diversity and are willing to work in both sectors even in the presence of wage differentials. Thus, the lower the elasticity $\rho$, the higher are the costs (measured in utility loss) of reallocating hours worked between sectors and the lower is intersectoral labour mobility.

The elasticity $\rho$ measures the size of imperfect labour market mobility and and will be translated into the empirical model. This modelling choice ought to be interpreted as a short-cut for more comprehensive models of labour market rigidities with the aim of keeping the analysis tractable and allowing for explicit analytical solutions and comparative statics. A wider interpretation of this specification is worker heterogeneity with regard to sector specific skills that cannot be acquired costlessly.[7]

The price of the tradable commodities is given by world markets and is normalised to unity. We let $P_t$ denote the relative price of non-tradable services to tradable commodities. Temporal budget constraints are given by

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

not change the results as long as preferences of the elderly are biased in favour of non-tradables.

[7] Alternatively, imperfect labour mobility can be modeled as sectoral labour adjustment costs as in Craighead (2009). However, price effects are not present in steady state in this case, so that adjustment costs cannot account for long-run relations.

and

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

$W_t^T$ and $W_t^N$ label wages in the two sectors taken as given and $S_t$ denotes household savings that are invested on international capital markets. 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 $(1+R^{*})/\pi_t$. We abstract from a pay-as-you-go pension system. The first order conditions of the household maximization problem read

$$\frac{C_{t+1}^{N}}{C_t^{T}} = \frac{\beta(1+R^{*})}{P_{t+1}} \tag{5}$$

$$\frac{L_t^{T}}{L_t^{N}} = \left(\frac{W_t^{T}}{W_t^{N}}\right)^{\rho}. \tag{6}$$

Euler equation (5) shows that the optimal ratio of non-tradables to tradables depends negatively on the relative price of the two. The OADR $\pi_t$ does not appear explicitly in this equation due to the assumption of perfect annuity markets. Condition (6) states that hours worked in both sectors depend on the wage ratio $W_t^T/W_t^N$ and the elasticity of substitution $\rho$.[8]

The tradable goods sector provides a homogeneous consumption good that is produced using labour and physical capital. The production function is

$$Y_t^{T} = F\left(A_t^{T}, K_t, L_t^{T}\right) = A_t^{T}\left(K_t\right)^{\alpha}\left(L_t^{T}\right)^{1-\alpha}, \tag{7}$$

where $A_t^T$ is productivity in the tradable sector and $K_t$ stands for the physical capital stock. Firms borrow $K_t$ on international capital markets and it is assumed that the full stock depreciates within one period. Profit maximization and perfect competition among firms imply

$$\frac{K_t}{L_t^{T}} = \left(\frac{R^{*}}{A_t^{T}\alpha}\right)^{-\frac{1}{1-\alpha}} \tag{8}$$

and

$$W_t^{T} = (1-\alpha)A_t^{T}\left(\frac{K_t}{L_t^{T}}\right)^{\alpha}\left(=W^{T}\right). \tag{9}$$

In this small open economy setting, the optimality conditions imply that the capital intensity is tied down by the world interest rate for a given level of productivity.

[8] All derivations of the model are relegated to an appendix, which is available on request.

Then, the wage in the tradable sector can be considered as constant.

The non-tradable sector provides a service good and employs raw labour $L_t^N$ as the sole input in production. We assume a linear production technology

$$Y_t^N = F\left(A_t^N, L_t^N\right) = A_t^N \cdot L_t^N, \tag{10}$$

where $A_t^N$ is productivity in that sector.[9] The corresponding first order condition reads

$$W_t^N = A_t^N \cdot P_t, \tag{11}$$

which relates the sectoral wage with the relative price and the sector-specific productivity parameter.

The equilibrium is defined taking the interest rate $R^*$ and the price of tradable goods as given by world markets. Households choose consumption $C_t^T$ and $C_t^N$ and sectoral labour supply $L_t^T$ and $L_t^N$, taking prices as given. Firms also operate as price takers. They choose labour demand in both sectors and capital demand in the tradable sector. Labour markets clear every period. Savings of households are fully invested at international capital markets and firms borrow all capital from abroad. The market clearing condition of the non-tradable sector is given by

$$Y_t^N = \pi_t C_t^N. \tag{12}$$

In the tradable goods sector, demand does not need to be met by domestic supply as $C_t^T$ can be traded on international markets.

Increasing the old-age dependency ratio $\pi_t$ leads to higher demand for non-tradable services $C_t^N$. In a model with perfect factor mobility this would be fully met by higher supply: labour would move into the service sector due to the positive wage pressure induced by higher demand. This would raise production of services until wages are equal in both sectors again, leaving relative prices unchanged. In contrast, with imperfect labour mobility, higher demand in the non-tradable service sector has a positive relative price effect because labour reallocation is not complete. Increased demand is only partly met by changes in supply and partly by an increasing relative price of non-tradables.

In order to analyse the impact of ageing on the relative price of non-tradables within our model, we proceed as follows: By means of market clearing condition (12), the demand and the production function of the non-tradable sector, we derive

[9]Extending the set of production inputs of the service sector to labour and capital merely complicates the analysis but does not alter our main results.

an implicit function of the relative price:

$$\frac{A_t^N P_t}{W_t^T}\left(1+\left(\frac{A_t^N P_t}{W_t^T}\right)^{1+\rho}\right)^{-\frac{1}{1+\rho}}-\frac{1+\beta\pi_t}{\beta\pi_t\left(1+R^*\right)}\left(1+\left(\frac{W_t^T}{A_t^N P_t}\right)^{1+\rho}\right)^{-\frac{2+\rho}{1+\rho}}=0 \quad (13)$$

Applying the implicit function theorem, we obtain the following observation:

**Observation 1** *The effect of ageing on the relative price of non-tradables is positive:*

$$\frac{\partial P_t}{\partial \pi_t}>0. \quad (14)$$

Furthermore, the relative price of non-tradables is also positively related to relative productivity $A_t^T/A_t^N$, which corresponds to the standard Balassa-Samuelson effect.

To analyse the importance of imperfect substitutability of labour between sectors for the impact of demographic change on relative prices, we present a calibrated version of the model in Table 1. Within the scope of the calibration, we increase the OADR from $\pi_t = 0.15$ to $\pi_t = 0.24$, corresponding to the increase in sample averages of the OADR between 1970 and 2009 in our data set (see Section 5). The table gives the change for relative prices and labour supply in each sector for a benchmark case with $\rho = 1$, an estimation taken from Horvath (2000), and for a case of higher intersectoral labour substitutability, assuming $\rho = 5$.

Table 1: CALIBRATED ECONOMY

| | Low Elasticity $\rho = 1$ | | High Elasticity $\rho = 5$ | |
|---|---|---|---|---|
| | $\pi = 0.15$ | $\pi = 0.24$ | $\pi = 0.15$ | $\pi = 0.24$ |
| Equilibrium price | 0.17 | 0.22 | 0.26 | 0.28 |
| Relative change | | 30% | | 8% |
| Labour Non-tradables sector, $L^N$ | 0.32 | 0.40 | 0.20 | 0.28 |
| Relative change | | 25% | | 40% |
| Labour Tradables sector, $L^T$ | 0.61 | 0.57 | 0.56 | 0.49 |
| Relative change | | −7% | | −13% |

Model parameters are chosen as follows: one period $t$ corresponds to 25 years, thus an annual interest rate of $R_{ann} = 2\%$ gives $1 + R = (1 + R_{ann})^{25}$. Assuming a discount rate of 2%, $\beta = 0.61$. Utility weight of leisure is set to $\phi = 0.5$. Relative productivity $A_t^T/A_t^N$ is set to 0.95, which is the average between 1970-2009 in our sample (see Section 5), and we set $A_t^N$. Benchmark labour elasticity $\rho = 1$ is taken from Horvath (2000). Capital share $\alpha$ is 0.3. $\pi_t$ is the average sample OADR in 1970 (15) and in 2009 (24).

According to our stylized model, the impact of an increase of the OADR on the relative price of non-tradables is sizable. In the benchmark specification, the change of $\pi_t$ from 0.15 to 0.24 leads to a 30 per cent increase of the relative price of non-tradable to tradable goods. Sectoral labour reallocates, as expected, to a limited extent in the direction of non-tradables. Increasing the intersectoral labour mobility by setting the $\rho = 5$, leads to a less pronounced increase in the relative prices of only 8 per cent, but to considerably stronger reallocations of labour. In the limiting case of $\rho \to \infty$ (not shown), which implies that workers devote all working time to the sector paying the highest wage, price effects completely vanish.

The calibration thus leads us to the following observation.

**Observation 2** *The relative price effect of ageing is the higher, the lower the labour mobility between the two production sectors.*

Observations (1) and (2) constitute the two hypothesis we are going to test in the empirical analysis.

# 4 Estimation Strategy

In order to obtain a testable linear econometric specification, we log-linearize the implicit function (13) that determines the relative sectoral price, assuming the economy to be in steady state. This procedure results in

$$\widehat{p}_t = \omega \widehat{\pi}_t + \frac{1}{1-\alpha}\widehat{a}_t^T - \widehat{a}_t^N \tag{15}$$

where $\omega = \left(\left(W^N/W^T\right)^{1+\rho} + 1\right) / \left(\left(1+\rho\right)\left(1+\beta\pi\right)\right)$. Lower case variables with hats ($\widehat{x}_t$) denote percentage deviations from steady state and variables without time index label values at steady state. The relative price thus depends on the old-age dependency ratio and the productivity differential between the two sectors, which corresponds to the standard Balassa-Samuelson effect. It can be shown that $\omega$ depends negatively on $\rho$.

Our first econometric estimation equation that follows from the log-linearized model is given by

$$p_{it} = \gamma_i + \gamma_1 oadr_{it} + \gamma_2 rpr_{it} + u_{it}, \tag{16}$$

where the sub-indices denote country $i$ and time period $t$ respectively. This specification focuses on the direct effect of ageing on relative prices and aims at testing the hypothesis stated in Observation 1. Thus, the covariate of main interest is $oadr_{it}$, which measures the old-age dependency ratio. According to our theory, its coefficient should possess a positive sign. To allow for an interpretation as (semi-)elasticities, variables are, when sensible, given as natural logarithms. So, $p_{it}$ denotes the log of a measure of the relative price of non-tradables and $rpr_{it}$ is the log of productivity in the tradables relative to the non-tradables sector. Theory predicts its coefficient to be positive and about unity, though this cannot always be confirmed in applied work.

Our second hypothesis, deduced from Observation 2, claims that imperfect labour mobility leads to higher price effects of ageing. This is tested by means of the following econometric specification:

$$p_{it} = \gamma_i + \gamma_1 oadr_{it} + \gamma_2 (oadr_{it} \cdot lri_{i(t)}) + \gamma_3 rpr_{it} + u_{it}, \tag{17}$$

where $lri$ denotes a measure of labour market rigidities, which is considered as the empirical counterpart of $\rho$. $lri$ is interacted with the OADR in order to assess the extent to which the effect of ageing changes with varying degrees of labour immobility. We apply different types of indices that are either fixed per country ($lri_i$) or vary over time ($lri_{it}$).

The two benchmark specifications are subsequently augmented with other variables discussed in the literature to control for further potential effects on relative prices. First, we add GDP per capita ($gdp_{it}$) to control for effects of factor-endowments in the spirit of Bhagwati (1984). Moreover, $gdp_{it}$ ought to be capable of capturing demand-side effects due to non-homothetic preferences, which consider non-tradable services as luxuries and tradable commodities as necessities – an approach proposed by Bergstrand (1991). In the presence of such effects, the coefficient of $gdp_{it}$ is expected to be positive. De Gregorio et al. (1994) propose government spending relative to GDP ($gov_{it}$) to control for further demand effects, since public expenditures are known to be biased towards non-tradables. Its coefficient is hence also anticipated to be positive. Lane and Milesi-Ferretti (2004) deal with wealth effects of net foreign asset (NFA) positions on real exchange rates. According to their argument, an increase in NFA 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 also augment our specifications with a variable on net foreign assets relative to GDP ($nfa_{it}$).

The econometric models all allow for country-specific intercepts $\gamma_i$ to pick up individual fixed effects. Following Pesaran (2006), $u_{it}$ represents an error-term of multi-factorial structure, given by

$$u_{it} = \boldsymbol{\delta}_i'\mathbf{f}_t + \varepsilon_{it} \tag{18}$$

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. Examples for $\mathbf{f}_t$ in case of economic macro variables are common business cycles as due to oil price shocks, the world financial crisis, or the effects of globalisation. In case of demographic variables, one may think of changes in working environments, habits, or medical innovations that increase longevity or the contraceptive pill. Meanwhile, $\varepsilon_{it}$ denotes an idiosyncratic error term.

Furthermore, the covariates of Models (16) and (17) can also be correlated with the same unobserved factors $\mathbf{f}_t$ as $u_{it}$ :

$$x_{it} = a_i + \boldsymbol{\eta}_i'\mathbf{f}_t + v_{it} \tag{19}$$

Here, $x_{it}$ denotes arbitrary RHS-variable of Models (16) and (17) that are 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 $v_{it}$.

The potential presence of $\mathbf{f}_t$ in both, (18) and (19) is responsible for the issue of cross-sectional correlation that typically arises in macro-applications. Disregarding $\mathbf{f}_t$ can bias standard errors of conventional estimators seriously and may at worst revert outcomes of empirical investigations, as for instance in O'Connell (1998) in the context of tests for purchasing power parity. An approach to remedy the problem is to apply the class of Common Correlated Effects (CCE) estimators developed by Pesaran (2006). These estimators are practically computed as ordinary least squares regressions augmented with cross-section averages of the dependent and all independent variables. Two different variants of the estimator exist – a pooled version that postulates homogeneous slope coefficients (called CCEP) and a mean-group version (called CCEMG) that is based on the average of individual-specific CCE estimates. The advantage of the latter is that it accounts for parameter heterogeneity, which comes at the cost of efficiency losses. Pesaran (2006) notes that the CCEP estimator is shown to perform better than the mean-group variant in small samples. Kapetanios et al. (2011) show that these estimators are also consistent in presence of unit roots in the unobservable factors. Monte Carlo studies by Kapetanios et al. (2011)

and by Coakley et al. (2006) further demonstrate the superiority of the CCEP estimator over other commonly used ones, even in small samples as ours. Thus, it is the estimator of our choice in this study.

Table 2: Sample Overview

| Country | Abbrev. | Coverage | Country | Abbrev. | Coverage |
|---|---|---|---|---|---|
| Austria | AUT | 1976-2009 | South Korea | KOR | 1970-2009 |
| 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 |

N=547 usable observations in benchmark models.

# 5 Data

## Data description

The empirical investigation is based upon a new constructed data set, consisting of a panel 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 547 usable observations in the benchmark models. The choice of countries is restricted by the availability of sufficiently detailed data on sectoral prices over sufficiently long time horizons. An overview of the sample along with some descriptive statistics on key variables is given in Table 2. A list of all variables used throughout the text can be found in Table 3.

All data stem from publicly available sources. For some variables standard data sets such as the OECD STAN data base or the Penn World Tables are used, for others we rely on data provided by other researchers. Details on the sources and regarding the construction of all variables are shifted to Appendix A.1. The relative price of non-tradable goods is constructed as the quotient of price indices of the non-tradables and the tradables sector. Vice versa, $rpr_{it}$ refers to the productivity ratio of the tradables relative to the non-tradables sector. To avoid measurement errors, productivities are always measured as labour, not as total factor productivi-

Table 3: List of Variables

| Variable | N | Description (Unit) |
|---|---|---|
| $p_{it}$ | 547 | Log of relative price of non-tradable to tradable sector (index) |
| $p_{it}^{var}$ | 444 | - variant: health care sector relative to tradables (index) |
| $oadr_{it}$ | 547 | Old-age dependency ratio (%) |
| $otp_{it}$ | 547 | Old over total population (%) |
| $lri_i$ | 547 | Labour market rigidity index, [0,1] |
| $lri_{it}^{var}$ | 482 | - variant: LAMRIG [0,3.5] |
| $rpr_{it}$ | 547 | Log of relative productivity of tradable to non-tradable sector (index) |
| $rpr_{it}^{var}$ | 405 | - variant: tradables relative to health care sector (index) |
| $gdp_{it}$ | 547 | Log of GDP per capita (2005 Int$) |
| $gov_{it}$ | 547 | Government consumption (% of GDP) |
| $nfa_{it}$ | 546 | Nef foreign assets (% of GDP) |

N reports number of available observations. 'Variant' denotes alternative measures of the corresponding variables used for sensitivity analyses.

ties (TFP).[10] The relative price of the health care sector $p_{it}^{var}$, used in the sensitivity analysis, is the ratio of the health price index and the price index of the tradable sector. The corresponding relative productivity measure $rpr_{it}^{var}$ is constructed in an analogous manner. The old-age dependency ratio, $oadr_{it}$, is defined as population aged older than 65 divided by population in working age (15-64). Similarly, $otp_{it}$ measures the amount of old people (65+) relative to total population. To give an impression of the magnitude and evolution over time of the data, we present some summary statistics in Table 4. 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, $lri_i$, in the benchmark models is taken from Botero et al. (2004). It is defined as the average of four other indices, namely (1) alternative employment contracts, (2) cost of increasing hours worked, (3) cost of firing workers, and (4) dismissal procedures. This composite index can attain values between zero and one, where higher values represent larger rigidities. Table 5 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. Campos and Nugent (2012) extend the index by Botero et al. (2004) to an even more comprehensive measure, the Labor

[10] For a broad discussion of this issue, the reader may be referred to Canzoneri et al. (1999).

Table 4: Summary Statistics

| Variable | 1970 | | 2009 | | Average |
|---|---|---|---|---|---|
| | Mean | Std. Dev. | Mean | Std. Dev. | Annual Change[a] |
| $\exp(p_{it})$ | 63.80 | 17.28 | 121.86 | 18.82 | 1.86 |
| $\exp(p_{it}^{var})$ | 48.43 | 11.28 | 135.58 | 30.14 | 3.07 |
| $oadr_{it}$ | 15.04 | 5.25 | 23.51 | 4.30 | 1.06 |
| $otp_{it}$ | 9.56 | 3.38 | 15.77 | 2.63 | 1.23 |
| $lri_i$ | 0.53 | 0.22 | 0.58 | 0.17 | – |
| $lri_{it}^{var}$ | 1.05 | 0.68 | 1.61[b] | 0.66[b] | 4.12 |
| $\exp(rpr_{it})$ | 52.86 | 21.73 | 148.75 | 42.01 | 2.65 |
| $\exp(rpr_{it}^{var})$ | 55.07 | 14.22 | 224.10 | 85.37 | 3.83 |
| $\exp(gdp_{it})$ | 14242.52 | 4943.40 | 34650.65 | 7425.53 | 2.31 |
| $gov_{it}$ | 7.36 | 1.35 | 7.20 | 1.18 | -0.01 |
| $nfa_{it}$ | -1.94 | 34.85 | -1.39 | 48.94 | 0.12 |

[a]Cross-sectional mean of average annual growth rates in per cent. [b]Mean and Std. Dev. in 2004 instead of 2009 due to data limitations.

Table 5: LRI per Country

| Country | $lri_i$ | Mean($lri_{it}^{var}$) | Country | $lri_i$ | Mean($lri_{it}^{var}$) |
|---|---|---|---|---|---|
| Austria | 0.50 | 1.40 | South Korea | 0.45 | 1.43 |
| Belgium | 0.51 | 1.49 | Netherlands | 0.73 | 2.18 |
| Canada | 0.26 | 0.48 | Norway | 0.69 | 2.11 |
| Denmark | 0.57 | 1.85 | Portugal | 0.81 | 2.39 |
| Finland | 0.74 | 1.85 | Spain | 0.74 | 2.78 |
| France | 0.74 | 1.84 | United Kingdom | 0.28 | 0.83 |
| Italy | 0.65 | 1.94 | United States | 0.22 | 0.42 |
| Japan | 0.16 | 0.54 | *Full Sample (avg.)* | 0.54 | 1.57 |

$lri_i$: see Botero et al. (2004). $lri_{it}^{var}$: see Campos and Nugent (2012).

Market Legislation Rigidity (LAMRIG) index. They extend the coverage to a higher number of countries and, more importantly, they construct a time-*variant* index (5-year averages between 1960 and 2004), which we denote as $lri_{it}^{var}$ and use for a robustness check. This alternative index is on a scale from 0 (least rigid) to 3.5 (most rigid). As Table 5 reveals, means per country of the LAMRIG index yield a similar ranking as the index by Botero et al. (2004).

## Non-stationarity and cross-sectional dependence

Before we move on to regression analysis, the data is tested for cross-sectional correlation and trend behaviour. To address the first issue, Table 6 presents average (absolute) cross-section correlation coefficients and results of Pesaran's (2004) $CD_P$

Table 6: Cross-section Dependence Tests

| | $CD_P$ | $avg.\left(\rho_{ij}\right)$ | $avg.\left(\lvert\rho_{ij}\rvert\right)$ |
|---|---|---|---|
| $p_{it}$ | 40.58*** | 0.809 | 0.844 |
| $p_{it}^{var}$ | 44.77*** | 0.894 | 0.894 |
| $oadr_{it}$ | 23.75*** | 0.472 | 0.611 |
| $otp_{it}$ | 31.03*** | 0.623 | 0.729 |
| $rpr_{it}$ | 43.13*** | 0.954 | 0.954 |
| $rpr_{it}^{var}$ | 43.15*** | 0.953 | 0.953 |
| $gdp_{it}$ | 43.76*** | 0.969 | 0.969 |
| $gov_{it}$ | 14.44*** | 0.324 | 0.472 |
| $nfa_{it}$ | 0.85 | 0.026 | 0.485 |

$CD_P$ denotes Pesaran (2004) cross-section dependence test statistic. All values but for $nfa_{it}$ are significant at 1 % (***) level. $avg.\left(\rho_{ij}\right)$ and $avg.\left(\lvert\rho_{ij}\rvert\right)$ denote average and average absolute cross-section correlation coefficients.

test statistic, which is $N(0,1)-$ distributed under the null hypothesis of cross-section independence. The $CD_P$ statistics are highly significant for all variables but NFA and the computed average correlation coefficients reveal strong correlations for all variables, but government consumption and NFA. Altogether, the results leave no doubt that cross-section correlation is indeed a problem in this data set.

Table 7: CIPS Panel Unit Root Tests

| | (a) intercept: | | | (b) intercept+trend: | | |
|---|---|---|---|---|---|---|
| | CADF(0) | CADF(1) | CADF(2) | CADF(0) | CADF(1) | CADF(2) |
| $p_{it}$ | 1.269 | 0.165 | 0.252 | 2.962 | 1.959 | 2.660 |
| $p_{it}^{var}$ | 0.512 | $-2.170^{**}_{IC}$ | -1.026 | 2.455 | $-0.058_{IC}$ | 1.024 |
| $oadr_{it}$ | 4.731 | -6.077*** | $5.073_{IC}$ | 6.680 | -5.931*** | $6.024_{IC}$ |
| $otp_{it}$ | 3.412 | -5.644*** | $4.537_{IC}$ | 4.911 | -7.373*** | $4.411_{IC}$ |
| $rpr_{it}$ | 0.456 | 0.321 | 1.162 | 2.881 | 2.284 | 4.021 |
| $rpr_{it}^{var}$ | 2.412 | 1.538 | 0.809 | 4.572 | 3.835 | 2.967 |
| $gdp_{it}$ | $-0.367_{IC}$ | -1.411* | -0.931 | 0.742 | 0.475 | 0.918 |
| $gov_{it}$ | 0.661 | 0.014 | 0.124 | 1.513 | 1.127 | 1.410 |
| $nfa_{it}$ | 6.244 | 5.511 | 5.832 | 2.240 | 0.920 | 1.536 |

Results of CIPS panel unit root tests using different lag lengths (0)-(2) without and with trend term. Asterisks indicate rejection of the null hypothesis of a unit root at 10%(*), 5% (**) and 1% (***). Subscript $IC$ indicates the model specification chosen by information criteria.

According to Maddala and Wu (1999), panel unit root tests that neglect cross-sectional correlation tend to have non-negligible size distortions. Thus, to test the order of integration of our data, we apply the second-generation CIPS panel unit root test, as advanced by Pesaran (2007), that accounts for that issue. In the spirit of Im et al. (2003), this test is based on standard augmented Dickey-Fuller regressions on the individual time series, but which are extended by cross-section averages (therefore also called CADF) both in lagged levels and first-differences of the variable in question. We present results for two variants of the test. The first one, which is presented in Panel (A) of Table 7, goes with an intercept in the CADF regressions, in order to test whether the series follow a random walk (with drift). The second one in Panel (B) includes both an intercept and a trend term to the regressions. Under the alternative hypothesis of this variant, the series are considered as trend-stationary processes. Furthermore, in order to control for serial correlation, the test statistics are provided for different lag augmentations, starting with zero and climbing up to two lags.

In the majority of cases, results in Table 7 reveal that variables can be marked as non-stationary, regardless of the inclusion of the trend term. In a few cases, outcomes are sensitive to the chosen lag length though. In these cases we use Akaike's and Schwarz's information criteria to determine the optimal model specification. In Panel (A), the CADF-statistic of $p_{it}^{var}$ appears to be significant if one lag is included, indicating that at least some individual time series are stationary. Both information criteria also select this lag length. However, in Panel (B), the null hypothesis of non-stationarity cannot be rejected. Noting that we lose about a fifth of the time series observations in case of $p_{it}^{var}$, this ambiguity may also be due to a type I error. Further inconclusiveness occurs in case of the two demographic variables and GDP per capita. For all three variables the information criteria choose non-stationary specifications to be optimal. Thus, we treat these as non-stationary also.

In order to assure that all series are integrated of the same order, we repeat the exercise for all variables in first differences. Results, which can be found in Appendix A.1, confirm that all series are difference-stationary. Accordingly, we treat all variables as integrated of order one.

# 6 Results

## Benchmark specification

We present estimates of the Models (16) and (17) using varying controls and estimation methods to assess our two hypotheses. For all regressions we provide further statistics on cross-section correlation and stationarity of the residuals to evaluate the validity of the findings.

Our main results on the first hypothesis, cf. Observation (1), that population ageing leads to higher relative prices of non-tradable goods are given in Table 8. All regressions in this table are conducted using the CCEP estimator. The benchmark specification (I) of Model (16) includes only the old-age dependency ratio and relative productivity. This is augmented consecutively by $gov_{it}$, $gdp_{it}$ and $nfa_{it}$ in (II) to (V). In all variants, $oadr_{it}$ enters the regressions significantly. The coefficients are estimated as semi-elasticities and can therefore be interpreted that a rise of the old-age dependency ratio by one percentage point increases the relative price of non-tradables by 0.6 to 1.3 per cent. As average old-age dependency ratios increase from about 15 to about 24 within the period covered by our sample, this 10 percentage point change implies an increase of relative prices of 6 to 13 per cent over the 40 years from 1970 to 2009 – an effect that seems reasonable and of non-negligible magnitude. The results are hence supportive to the first hypothesis.

Most of the control variables also enter statistically significant with coefficients roughly in line with theory. As relative productivity is seen as the main determinant of relative price movements throughout the literature, it is the only control variable that is included in all five specifications. Coefficients on $rpr_{it}$ suggest that a one per cent increase results in about 0.7 per cent higher relative prices. The standard Balassa-Samuelson model as well as (15) predict coefficients of about unity for relative TFP, but given that relative labour productivity is used to proxy TFP, these estimates are satisfactory. The fact that the effect of $gdp_{it}$ is positive and significant indicates the presence of further supply or demand effects as discussed in Section 4. However, government expenditures, which ought to capture further demand effects, show up wrong-signed or insignificantly. This seemingly odd finding may be explained by the work of Galstyan and Lane (2009), who note that parts of public spending, namely government investment, can also act deflating on relative prices. This might counteract the demand effects of higher non-tradable government consumption. Finally, in specification (V) net foreign assets that aim at controlling for all kinds of international wealth effects does not point in the direction predicted by theory either and its coefficient is estimated to be very small. Nevertheless, it is remarkable how stable the effect of ageing is estimated, independent of the choice of controls.

As residual diagnostics we test for cross-sectional dependence and stationary. Though, average cross-section correlation coefficients $\rho_{ij}$ of the residuals are close to zero, $CD_P$ test statistics in all cases still imply that the null hypothesis of cross-section independence is rejected. To test stationarity of the residuals, we thus apply

Table 8: CCEP Estimation

| $p_{it}$ | (I) | (II) | (III) | (IV) | (V) |
|---|---|---|---|---|---|
| $oadr_{it}$ | 0.00598** | 0.0104*** | 0.00735* | 0.0133*** | 0.00633* |
| | (0.00316) | (0.00343) | (0.00449) | (0.00486) | (0.00357) |
| $rpr_{it}$ | 0.689*** | 0.744*** | 0.599*** | 0.610*** | 0.595*** |
| | (0.0551) | (0.0495) | (0.0608) | (0.0540) | (0.0657) |
| $gov_{it}$ | | -0.0193*** | | 0.0103 | |
| | | (0.00664) | | (0.00838) | |
| $gdp_{it}$ | | | 0.225*** | 0.350*** | |
| | | | (0.0615) | (0.0879) | |
| $nfa_{it}$ | | | | | -0.000817*** |
| | | | | | (0.000265) |
| Residual cross-sectional dependence | | | | | |
| $CD_P$ | -2.09** | -3.30*** | -2.16** | -2.66*** | -2.48** |
| $avg\left(\rho_{ij}\right)$ | -0.035 | -0.054 | -0.036 | -0.044 | -0.041 |
| Residual stationarity | | | | | |
| $CADF(0)$ | -2.547** | -4.880* | -3.693** | -5.333** | -5.172* |
| $CADF(1)$ | -2.823+ | -4.496 | -3.736* | -5.200* | -5.697 |
| $CADF(2)$ | -1.731+ | -2.648+ | -1.858* | -3.322* | -3.620 |
| $N$ | 547 | 547 | 547 | 547 | 546 |

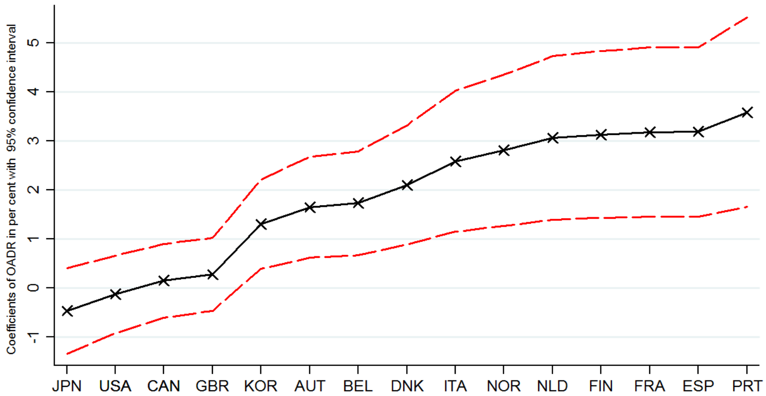
CCEP estimation results for different variants of Model (16). Standard errors in parentheses. Asterisks mark significance at 10% (*) 5% (**), 1% (***). Country dummies are included. Residual diagnostics: cross-section dependence test by Pesaran (2004), $CD_P$, and average cross-section correlation coefficient $\left(\rho_{ij}\right)$. Residual stationarity is tested with the Pesaran (2007) $CADF(j)$ test statistic with $j$ lags. Critical values are taken from a bootstrapped distribution. Asterisks denote a rejection of the Null of non-stationarity at 15% (+), 10% (*) and 5% (**).

the CIPS test again to account for these dependencies as for instance in Holly et al. (2010). Since the critical values for this statistic have not yet been tabulated for residual testing, we have tested for the significance of the CIPS statistics using a bootstrap approach.[11] Following Fachin (2007) we construct pseudo-data by applying the Continuous-Path Block Bootstrap (CPBB) method developed by Paparoditis and Politis (2000, 2003). The method is explicitely designed to preserve non-stationarity and the cross-sectional dependence of the data. As suggested by Paparoditis and Politis (2003), we implement the block bootstrap for our unbal-

[11] Banerjee and Carrion-i Silvestre (2011) only provide critical values for up to two covariates using Monte-Carlo techniques.

anced panel assuming a fixed block length of 10 percent of the overall observation period. The critical values are computed as follows: Pseudo-data is constructed by use of the original panel data set. We then run the CCEP estimations and compute the residual test statistic of the CIPS test. We redraw the bootstrap 500 times and compute the distribution of the CIPS test statistics. The residuals are considered as stationary, if the CIPS test statistic falls outside a fixed confidence interval of the bootstrapped distribution. Moscone and Tosetti (2010) recently applied this method and assumed stationarity of the residuals if the statistic is outside the 25th and 75th percentiles of the bootstrapped distribution. We use the stricter bounds of 15th and 85th percentiles of the bootstrapped distribution (marked with a $^{+}$), as well as the 90th and 10th percentiles (marked with $^{*}$) and the 95th and the 5th percentiles (marked with $^{**}$) of the distribution. The lower panel of Table 8 shows the $CADF(j)$ statistics for lag orders $j = 0, 1, 2$. The test suggests that residuals of all specifications can be seen as integrated of order zero for most lag orders.

Figure 2: Interaction Effects - Visualisation




Coefficients and 95% confidence interval of oadr$_{it}$ evaluated at the lri$_i$-values of all countries in the sample (see Table 5) using Specification (IX), Table (9).

Table 9 and Figure 2 present results on Model (17), which aims at testing the second hypothesis deduced from Observation (2) that price effects of population ageing are higher, the lower labour mobility is. To this end, the specifications of Table 8 are augmented with interactions of $lri_i$ and $oadr_{it}$. To give the output a meaningful and easy-to-read interpretation, we reparameterise the regression models such that the coefficients on $oadr_{it}$ in Table 9 give the effect of $oadr_{it}$ on $p_{it}$ at the median of $lri_i$. Again, it can be seen that a rise in the old-age dependency ratio leads to significant increases of relative prices, no matter what controls are added to the regressions. At the median of $lri_i$, coefficients of $oadr_{it}$ are estimated to be between

Table 9: CCEP Estimation - Interaction Effects

| $p_{it}$ | (VI) | (VII) | (VIII) | (IX) | (X) |
|---|---|---|---|---|---|
| $oadr_{it}$ | 0.0127*** | 0.0192*** | 0.0100** | 0.0210*** | 0.0111*** |
| | (0.00371) | (0.00390) | (0.00520) | (0.00618) | (0.00417) |
| $oadr_{it} \cdot lri_i$ | 0.0719*** | 0.0867*** | 0.0383** | 0.0628*** | 0.0628*** |
| | (0.0133) | (0.0139) | (0.0166) | (0.0176) | (0.0143) |
| $rpr_{it}$ | 0.672*** | 0.721*** | 0.596*** | 0.597*** | 0.585*** |
| | (0.0507) | (0.0431) | (0.0588) | (0.0504) | (0.0618) |
| $gov_{it}$ | | -0.0170*** | | 0.00452 | |
| | | (0.00646) | | (0.00806) | |
| $gdp_{it}$ | | | 0.179*** | 0.274*** | |
| | | | (0.0622) | (0.0810) | |
| $nfa_{it}$ | | | | | -0.000816*** |
| | | | | | (0.000251) |
| Residual cross-sectional dependence | | | | | |
| $CD_P$ | -2.59*** | -3.32*** | -2.28** | -2.76*** | -2.83*** |
| $avg\left(\rho_{ij}\right)$ | -0.043 | -0.054 | -0.038 | -0.046 | -0.046 |
| Residual stationarity | | | | | |
| $CADF(0)$ | -2.195** | -5.059+ | -3.702** | -5.871+ | -4.544** |
| $CADF(1)$ | -2.872+ | -4.856 | -3.972* | -6.143 | -4.980 |
| $CADF(2)$ | -2.041 | -3.712 | -2.625+ | -4.403 | -3.840 |
| $N$ | 547 | 547 | 547 | 547 | 546 |

CCEP estimation results for different variants of Model (17). Standard errors in parentheses. Asterisks mark significance at 10% (*) 5% (**), 1% (***). Country dummies are included. Residual diagnostics: cross-section dependence test by Pesaran (2004), $CD_P$, and average cross-section correlation coefficient $avg\left(\rho_{ij}\right)$. Residual stationarity is tested with the Pesaran (2007) $CADF(j)$ test statistic with $j$ lags. Critical values are taken from a bootstrapped distribution. Asterisks denote a rejection of the Null of non-stationarity at 15% (+), 10% (*) and 5% (**).

one and two per cent. Patterns of magnitude and significance of the other coefficients as well as of residual diagnostics resemble those in Table 8. To visualise the effect of the old-age dependency ratio on relative prices at different levels of labour market rigidities, we present coefficients of $oadr_{it}$ exemplarily from specification (IX) for the $lri_i$-values of all countries in Figure 2. Conforming to the theoretical predictions, countries with the most flexible labour markets, appearing on the left side of the figure, undergo very small price effects of ageing, while countries with more rigid markets can be seen to experience larger effects. In particular, for $lri_i$-values up to

0.30 that are related with Anglo-American countries price effects are not estimated to be statistically different from zero. In case of $lri_i$-values about 0.70-0.80, which correspond to Southern European countries as France, Spain and Portugal, relative price effects rise up to over three per cent.[12] This is clear evidence of our second hypothesis of Observation (2) and supports the validity of the proposed transmission via imperfect labour market mobility.

Next, we provide results on a series of sensitivity analyses to demonstrate the robustness of our findings. First, we substitute all main variables – the relative price, demography, and the labour rigidity index – by alternative measures. Second, we redo the estimations using different methods in place of CCEP.

## Variations in Dependent and Independent Variables

Table 10: Variations in Dependent and Independent Variables

| Variation 1: $p_{it}^{var}$ instead of $p_{it}$ | (I A) | (II A) | (III A) | (IV A) | (V A) |
|---|---|---|---|---|---|
| $oadr_{it}$ | 0.0136** | 0.0220*** | 0.0140* | 0.0270*** | 0.0130** |
| | (0.00558) | (0.00639) | (0.00797) | (0.00981) | (0.00638) |
| $N$ | 405 | 405 | 405 | 405 | 405 |
| Variation 2: $otp_{it}$ instead of $oadr_{it}$ | (I B) | (II B) | (III B) | (IV B) | (V B) |
| $otp_{it}$ | 0.0102* | 0.0174*** | 0.00878 | 0.0158** | 0.0108* |
| | (0.00547) | (0.00603) | (0.00649) | (0.00774) | (0.00607) |
| $N$ | 547 | 547 | 547 | 547 | 546 |
| Variation 3: $lri_{it}^{var}$ instead of $lri_i$ | (VI A) | (VII A) | (VIII A) | (IX A) | (X A) |
| $oadr_{it}$ | 0.0133*** | 0.0161*** | 0.0139** | 0.0209*** | 0.0128*** |
| | (0.0047) | (0.0052) | (0.0061) | (0.0069) | (0.0045) |
| $oadr_{it} \cdot lri_{it}^{var}$ | 0.0031*** | 0.0031*** | 0.0015** | 0.0014** | 0.0029*** |
| | (0.0008) | (0.0007) | (0.0007) | (0.0006) | (0.0007) |
| $N$ | 482 | 482 | 482 | 482 | 482 |

CCEP estimation results for Models (16) and (17) using alternative variables. Roman numerals refer to the same specifications as in Tables 8 and 9. Standard errors in parentheses. Asterisks mark significance at 10% (*) 5% (**), 1% (***). Complete set of results can be obtained on request.

Table 10 shows results on regressions in the specifications of Tables 8 and 9 with variations in $p_{it}$, $oadr_{it}$, and $lri_i$.[13] According to the observations of Section 2, the

[12] Of course, effects evolve similarly when estimated from one of the other model specifications.

[13] For the sake of brevity, we present coefficients of the demographic variables only.

share of health care expenditures of the elderly is on average about 30 per cent larger than for people in working age – health care thereby being the expenditure group with the largest changes over the life cycle. Thus, we conjecture that price changes in the health care sector should be extraordinary large. To test the hypothesis, we employ $p_{it}^{var}$ and $rpr_{it}^{var}$ instead of $p_{it}$ and $rpr_{it}$ in Variation 1, which are price and productivity indices of the health care sector relative to the tradable goods sector. Indeed, we find this to be the case with coefficients being about twice as large as in Table 8.[14] However, as opposed to the former results, tests on residual stationarity reveal that the null of non-stationarity can in general not be rejected in Variation 1. Part of the reason for this differing behaviour may be the loss of power due to the considerable reduction in sample size.

In the next experiment, Variation 2, we substitute $oadr_{it}$ with the alternative demographic measure $otp_{it}$. Coefficients are now slightly larger than in the default models, though in case of specification (III B) the effect shows up to insignificant. Overall, the effects of population ageing stay qualitatively the same.

Finally, we employ the alternative time-varying labour market rigidity index by Campos and Nugent (2012) in Variation 3. Using this alternative index, we are able to retain significant effects of comparable size in all cases. Evaluating the coefficients of $oadr_{it}$ at different points of $lri_{it}^{var}$ yields analogical outcomes as those shown in Figure 2. In both, Variation 2 and 3, tests on residuals are primarily able to reject the null of non-stationarity. Taken together, our main results stay valid with the use of alternatives for all key variables.

## Variations in Estimation Models

The first alternative estimation model that we consider is a panel dynamic OLS estimator with one lead and one lag, named DOLS(1,1). Table 14 of Appendix A.2 provides highly significant coefficients for OADR throughout all specifications with magnitudes somewhat larger than in the benchmark. The coefficients of the control variables also resemble those of Table 8. However, tests on the residuals reject cross-sectional independence in all cases and non-stationarity in case of specifications (I) to (III). Though it is known that coefficients can be estimated consistently in non-stationary panels, we cannot rely on inference in these cases. Thus, controlling for potentially non-stationary unobserved common factors as it is done by the CCE procedure indeed seems to help in obtaining stationary regression residuals.

[14] All results not shown in the text are available on request.

In order to tackle the issue of non-stationarity of our data in another way, we use data in first differences that is found to be stationary in Appendix A.1 and re-estimate effects by the CCEP and by a standard First Difference OLS (FDOLS) estimator. Results in Tables 15 and 16 show significant coefficients for the old-age dependency ratio with magnitudes of about one per cent, which resembles the output of data in levels closely. In terms of significance, the results for the control variables are similar as well, though the coefficients on relative productivity shrink to an unreasonable size of about 0.2 per cent. Once more, the null of cross-section independence of the $CD_P$ test has to be rejected in each case, with the FDOLS results slightly worse than those on FD-CCEP. Aside to the invariance of our main results to the usage of data in first differences, effects also remain in place when trend terms are included to Models (16) or (17). Hence, apart from some non-stationarity issues in case of DOLS, the alternative estimation models all support our main outcome of higher relative prices of non-tradables due to population ageing.

# 7 Conclusion

In this paper we analyse the impact of population ageing on the relative price of non-tradable services to tradable commodities. First, we illustrate by means of a simple OLG model with two production sectors how population ageing can affect relative prices. Imperfect labour market mobility is key for the transmission of demand effects on relative prices in our setup. In the empirical tests, we identify a statistically and economically significant relation of reasonable size between the old-age dependency ratio and the relative price of non-tradables. We find support for our hypothesis that labour market rigidities are responsible for the effects. 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. As trends in population ageing for the countries of our sample are forecasted yet to exacerbate in the upcoming decades, one can expect considerable price changes due to demographic change.

# A Appendix

## A.1 Data Set Construction and Variable Properties

An overview of all data sources is given in Table 12. In all cases but relative prices and productivities no further data transformations 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 its member states, where total value added is decomposed into nine standardised 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 per cent, a sector is marked as being tradable. Using more recent data, Bettendorf and Dewachter (2007) repeat this exercise and are able to confirm the original classifications. Thus, 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 11. Accordingly, five sectors, accountable for 65 per cent of total value added in the year 2000, are classified as non-tradable, the four remaining sectors, accounting for 35 per cent, as tradable. As one can see, all service sectors except for '*Transport, Storage and Communications*' that is accountable for only 6.7 per cent of total value added, are marked as non-tradable – thereby justifying the practice to talk about tradables as commodities and non-tradables as services synonymously.

Deflators are computed to yield a separate price index of non-tradable services and tradable commodities using the following formula:

$$P_j = \frac{\sum_{s=1}^{j} VALU_s}{\sum_{s=1}^{j} VALK_s} \text{ for } j = N, T \tag{20}$$

where $VALU$ 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 $P = P_N/P_T$, which is – after taking logs – employed in the regressions. Data on relative productivity also stems from the STAN database. As mentioned earlier, we use labour productivity as a proxy for TFP. First, productivity measures for both, the non-tradable service and the tradable commodities sector are calculated by dividing sectoral value added at

Table 11: 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 |

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).

constant prices ($VALK$) by sectoral total employment ($EMPN$):

$$SPR_j = \frac{VALK_j}{EMPN_j} \text{ for } j = N, T \tag{21}$$

Relative productivity as used in the regression analysis ($rpr_{it}$) is then constructed by dividing $SPR_T$ by $SPR_N$ and taking logs of the result. The alternative measures $p_{it}^{var}$ and $rpr_{it}^{var}$ that we use for purposes of sensitivity analysis are constructed in a similar fashion. The difference is that here the measures for the non-tradable goods sector are substituted by respective measures of the sector '*Health and social work*', which is a subsector of '*Community, social and personal services*'. Thereby, we obtain a price and a productivity index of health care relative to tradable goods.

Table 12: 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 ($p_{it}, p_{it}^{var}, rpr_{it}, rpr_{it}^{var}$) | OECD - Structural Analysis (STAN): http://stats.oecd.org/ | VALU, VALK, EMPN |
| Demographic indicators ($oadr_{it}$, $otp_{it}$) | World Bank - World Development Indicators (WDI): http://data.worldbank.org/ | Age dependency ratio, old. Population ages 65 and above |
| Labour market rigidity index ($lri_i$) | Botero et al. (2004): http://faculty.tuck.dartmouth.edu/rafael-laporta/research-publications | index_labor7a |
| Labour market rigidity index - variant ($lri_{it}^{var}$) | Campos and Nugent (2012): http://www.naurocampos.net/papers/LAMRIG.html | LAMRIG |
| GDP per capita ($gdp_{it}$), government consumption ($gov_{it}$) | Heston et al. (2012) - Penn World Table (PWT) 7.1: http://pwt.econ.upenn.edu/ | rgdpl, kg |
| Net Foreign Assets ($nfa_{it}$) | Lane and Milesi-Feretti (2007): http://www.philiplane.org/EWN.html | NFA/GDP |

In Section 5, we found all variables to be non-stationary when tested in levels. Here, we repeat unit root tests for data in first differences. The results can be found in Table 13. As expected, for all variables the null hypothesis of a unit root has to be rejected and hence all variables are marked as stationary. The only exemptions are the two demographic variables in case when both an intercept and a trend are included in the test equation. However, information criteria choose the lag lengths, where test values are closest to a rejection of the null hypothesis. Furthermore one may take into consideration that including the trend term can lead to a misspecification of the models. Thus, taken as a whole, we conclude that all variables are difference-stationary and integrated of order one.

Table 13: CIPS Panel Unit Root Tests - First Differences

| | (A) intercept: | | | (B) intercept+trend: | | |
|---|---|---|---|---|---|---|
| | CADF(0) | CADF(1) | CADF(2) | CADF(0) | CADF(1) | CADF(2) |
| $\Delta p_{it}$ | -12.227*** | -6.984*** | -3.797*** | -11.849*** | -6.396*** | -2.620*** |
| $\Delta p_{it}^{var}$ | -9.353*** | -4.581*** | -2.957*** | -9.016*** | -3.712*** | -1.573* |
| $\Delta oadr_{it}$ | 5.967 | $-2.860^{***}_{IC}$ | 4.313 | 10.653 | $-0.509_{IC}$ | 6.897 |
| $\Delta otp_{it}$ | 5.854 | $-4.308^{***}_{IC}$ | 2.229 | 11.499 | $0.205_{IC}$ | 6.285 |
| $\Delta rpr_{it}$ | -11.883*** | -6.844*** | -4.163*** | -11.611*** | -7.391*** | -4.714*** |
| $\Delta rpr_{it}^{var}$ | $-8.549^{***}_{IC}$ | -4.688*** | -0.658 | $-7.380^{***}_{IC}$ | -3.992*** | 0.345 |
| $\Delta gdp_{it}$ | -11.753*** | -6.811*** | -4.966*** | -10.527*** | -5.224*** | -3.169*** |
| $\Delta gov_{it}$ | -11.834*** | -7.007*** | -3.683*** | -10.062*** | -5.000*** | -1.437* |
| $\Delta nfa_{it}$ | -12.251*** | -6.081*** | -3.627*** | -11.900*** | -5.401*** | -2.685*** |

Results of CIPS panel unit root tests using different lag lengths (0)-(2) without and with trend term on data in first differences (denoted by $\Delta$). Asterisks indicate rejection of the null hypothesis of a unit root at 10%(*), 5% (**) and 1% (***). Subscript $IC$ indicates the model specification chosen by information criteria.

## A.2 Variations in Estimation Models

Table 14: DOLS(1,1) Estimation

| $p_{it}$ | (I) | (II) | (III) | (IV) | (V) |
|---|---|---|---|---|---|
| $oadr_{it}$ | 0.0123*** | 0.0177*** | 0.0115*** | 0.0194*** | 0.0110*** |
| | (0.00226) | (0.00273) | (0.00227) | (0.00272) | (0.00193) |
| $rpr_{it}$ | 0.550*** | 0.533*** | 0.576*** | 0.624*** | 0.583*** |
| | (0.0302) | (0.0301) | (0.0384) | (0.0387) | (0.0259) |
| $gov_{it}$ | | -0.0207*** | | -0.0346*** | |
| | | (0.00625) | | (0.00701) | |
| $gdp_{it}$ | | | -0.0510 | -0.162*** | |
| | | | (0.0356) | (0.0406) | |
| $nfa_{it}$ | | | | | -0.00213*** |
| | | | | | (0.000166) |
| Residual cross-sectional dependence | | | | | |
| $CD_P$ | -2.71*** | -2.36** | -2.74*** | -2.49** | -3.39*** |
| $avg\left(\rho_{ij}\right)$ | -0.047 | -0.040 | -0.048 | -0.043 | -0.058 |
| Residual stationarity | | | | | |
| $CADF\,(0)$ | 1.021 | 0.672 | 0.560 | 0.464 | -1.076** |
| $CADF\,(1)$ | 0.647 | 1.307 | 1.584 | 0.608 | -1.113** |
| $CADF\,(2)$ | 1.925 | 2.635 | 1.801 | 0.085** | -0.215* |
| $N$ | 502 | 502 | 502 | 502 | 501 |

DOLS(1,1) estimation results for different variants of Model (16). Standard errors in parentheses. Asterisks mark significance at 10% (*) 5% (**), 1% (***). Country dummies are included. Residual diagnostics: cross-section dependence test by Pesaran (2004), $CD_P$, and average cross-section correlation coefficient $avg\left(\rho_{ij}\right)$. Residual stationarity is tested with the Pesaran (2007) $CADF(j)$ test statistic with $j$ lags. Critical values are taken from a bootstrapped distribution. Asterisks denote a rejection of the Null of non-stationarity at 15% (+), 10% (*) and 5% (**).

Table 15: FD-CCEP Estimation

| $\Delta p_{it}$ | (I) | (II) | (III) | (IV) | (V) |
|---|---|---|---|---|---|
| $\Delta oadr_{it}$ | 0.0135** | 0.0139* | 0.0151** | 0.0141* | 0.0120* |
| | (0.00696) | (0.00726) | (0.00735) | (0.00791) | (0.00712) |
| $\Delta rpr_{it}$ | 0.262*** | 0.316*** | 0.268*** | 0.289*** | 0.255*** |
| | (0.0715) | (0.0797) | (0.0782) | (0.0788) | (0.0755) |
| $\Delta gov_{it}$ | | 0.00321 | | 0.0182** | |
| | | (0.00779) | | (0.00870) | |
| $\Delta gdp_{it}$ | | | 0.117 | 0.268*** | |
| | | | (0.0841) | (0.0939) | |
| $\Delta nfa_{it}$ | | | | | -0.000350** |
| | | | | | (0.000181) |
| Residual cross-sectional dependence | | | | | |
| $CD_P$ | -2.75*** | -2.87*** | -2.65*** | -2.74*** | -2.57*** |
| $avg\left(\rho_{ij}\right)$ | -0.046 | -0.048 | -0.044 | -0.045 | -0.043 |
| $N$ | 532 | 532 | 532 | 532 | 531 |

CCEP estimation results with data in first-differences for different variants of Model (16). Standard errors in parentheses. Asterisks mark significance at 10% (*) 5% (**), 1% (***). Country dummies are included. Residual diagnostics: cross-section dependence test by Pesaran (2004), $CD_P$, and average cross-section correlation coefficient $avg\left(\rho_{ij}\right)$

Table 16: FDOLS Estimation

| $\Delta p_{it}$ | (I) | (II) | (III) | (IV) | (V) |
|---|---|---|---|---|---|
| $\Delta oadr_{it}$ | 0.0133** | 0.0112* | 0.0129** | 0.0112* | 0.0129** |
| | (0.00619) | (0.00618) | (0.00621) | (0.00618) | (0.00621) |
| $\Delta rpr_{it}$ | 0.169*** | 0.216*** | 0.190*** | 0.199*** | 0.173*** |
| | (0.0533) | (0.0550) | (0.0573) | (0.0569) | (0.0535) |
| $\Delta gov_{it}$ | | 0.0199*** | | 0.0254*** | |
| | | (0.00641) | | (0.00806) | |
| $\Delta gdp_{it}$ | | | -0.0588 | 0.0823 | |
| | | | (0.0594) | (0.0741) | |
| $\Delta nfa_{it}$ | | | | | 0.0000886 |
| | | | | | (0.000189) |
| Residual cross-sectional dependence | | | | | |
| $CD_P$ | 4.04*** | 2.70*** | 3.65*** | 2.79*** | 4.08*** |
| $avg\left(\rho_{ij}\right)$ | 0.069 | 0.046 | 0.063 | 0.048 | 0.070 |
| $N$ | 532 | 532 | 532 | 532 | 531 |

FDOLS estimation results with data in first-differences for different variants of Model (16). Standard errors in parentheses. Asterisks mark significance at 10% (*) 5% (**), 1% (***). Country dummies are included. Residual diagnostics: cross-section dependence test by Pesaran (2004), $CD_P$, and average cross-section correlation coefficient $avg\left(\rho_{ij}\right)$

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