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Ageing and the Relative Price of Nontradeables

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Ageing and the relative price of nontradeables

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Abstract

In this paper we identify the effects of ageing on the relative price of nontradeables versus tradeables. We consider two cases. In a first specification, age effects only account for short-run dynamics. An alternative case allows for permanent age effects. Estimating the respective cases by means of an ECM on a panel of OECD countries we find significant effects of demographic composition on the relative prices, even after correcting for the standard explanatory variables. Simulations based on population projections of the UN show that ageing might substantially contribute to inflationary pressures in the near future.

Keywords*:* relative price of nontradeables; age-specific effects; productivity differentials

JEL classification*:* F41, E31, J11

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1 Introduction

Demographic factors have a significant impact on various macroeonomic variables (such as consumption, output growth and inflation). Various studies have identified and or corroborated these effects of demographics on the macroeconomy, see e.g. Fair and Dominguez (1991) and Lindh and Malmberg (2000). In this paper we assess the impact of demographic factors on the relative price of notradeables versus tradeables. We review the alternative channels identified in the literature and subsequently empirically estimate the impact of demographics on the relative price. We find that demographic composition has significant effects on the relative prices.

Following the long-standing theoretical and empirical literature[1], there is reason to believe that changes in the demographic structure may affect the relative prices. The literature distinguishes three structural determinants of the relative price of nontradeables (Bergstrand, 1991): productivity differentials (Balassa-Samuelson), relative factor endowments (Bhagwati) and relative demand shocks (Bergstrand).

First, according to the Balassa-Samuelson model, the relative price is explained by productivity differentials. Since productivity in the traded goods sector grows faster than in the non-traded goods sector, the relative price of nontradeables increases to match the nominal wage increases induced by the productivity increases. Unfortunately, the relation between the demographic structure and aggregate productivity is still unclear, let alone the impact on sectoral productivities (Cutler et al., 1990). As such it is a priori unclear to what extent demographic structure might affect productivity differentials between tradeables and nontradeables.

A second, supply-oriented hypothesis (Bhagwati) focuses on the role of relative factor endowments. As nontradeables are relatively labor intensive in production and tradeables relatively capital intensive, a larger relative endowment of capital will increase the relative price. This second type of determinants clearly identifies a supply side effect of ageing. In particular, following the life cycle theory, factor endowments have an age component. Older age groups are characterized by low labor participation and low savings (or even dissavings), compared to younger age groups. When labor gets relatively scarce due to the ageing of the labor force, the labor intensive sectors producing nontradeables are in particular affected. This second channel thus identifies a positive relation between the dependency ratio and the relative price of nontradeables.

A last theory (Bergstrand) points at the link with the relative demand structure. Under the assumption that nontradeables are luxuries in consumption and tradeables are necessities, income growth will cause a higher relative price. Next to income growth, one

[1]See Froot and Rogoff (1995) for an overview of this literature.

could argue that demographic structure may affect relative demand as well. As regards the demand hypothesis, older consumers seem to spend a larger fraction of their budget on nontradeables (like health services), irrespective of favorable income changes.[2] Aggregate demand for nontradeables growing relatively faster than the demand for tradeables contributes to upward pressures on the relative price of nontradeables.

The above channels identify both demand and supply factors through which demographics may affect relative prices. Also, some factors may have a more permanent effect (typically the supply factors) while others (demand factors) are considered more as temporary factors. In order to accommodate both temporary and permanent factors we use an ECM approach in estimating the model. More in particular we proceed in two steps. In a first step we adhere to the standard view in which demographic effects are considered as temporary. We thus model relative prices using the standard Balassa-Samuelson cointegrating relation and estimate the temporary demographic effects. Subsequently, we allow the demographics to have a permanent effect by analyzing the effects between demographics and (relative) productivity trends in the tradeable and nontradeable sector. We find, especially in the latter case, significant effects of ageing on relative prices. The potential consequences of demographic developments and more in particular ageing are illustrated by simulating the estimated relations with population projections taken from the United Nations.

The remainder of the paper is structured as follows. The next section briefly discusses the determination of the relative price in a simplified theoretical framework. The construction of the panel dataset is explained in Section 3. We first estimate an ECM in which age effects only matter for short-run dynamics. After the panel cointegration test is passed, results are presented in Section 4. The effects of the future demographic transition are simulated for this case in Section 5. Section 6 analyzes the long-run effects of demographics by establishing empirically a link between relative productivity and the demographic (age) structure. Subsequently, we estimate (and simulate) an alternative ECM with permanent age effects. The last Section summarizes and mentions some policy implications.

[2] Hobijn and Lagakos (2003) show that the spending pattern of the typical retiree differs significantly from that of the typical worker for the US. In 2001, the budget share of a retiree for housing and medicare was 8 and 5 percent points higher than for an urban worker, respectively. As a result, a retiree faced in fact a higher inflation rate in each year over the 1984-2001 period. See Börsch-Supan (2001) for evidence for Germany.

2 Theory

This section provides a brief overview of theoretical insights in the determination of the relative price, as explained by Rogoff (1992). It aims to illustrate the role of demand changes induced by ageing on relative prices using a simplified model. Obviously, a detailed analysis of age-specific effects needs a full-fledged model with overlapping generations.[3] This is beyond the scope of this empirical paper.

2.1 The model

We consider a representative agent model of a small, open economy with a nontradeable (N) and a tradeable (T) sector. Production in both sectors uses labor and capital with a Cobb-Douglas technology:

$$Y_x = \theta_x L_x^{\alpha_x} K_x^{1-\alpha_x} \qquad x = \{N, T\} \tag{1}$$

where Y_x, L_x and K_x denote output, labor and capital in sector x, respectively. The lifetime utility function of the representative agent is specified as:

$$U_t = \sum\nolimits_{s=t}^{\infty} (1+r)^{t-s} \ln \left(C_{N,s}^{\phi} C_{T,s}^{1-\phi} \right) \tag{2}$$

where C_N and C_T denote the consumption of nontradeables and tradeables, respectively. The rate of time preference equals the world interest rate r and the intratemporal, as well as the intertemporal substitution elasticities are assumed equal to one. Investment goods produced by sector x are denoted by I_x.[4] Government expenditures G_x are exogenous and financed by lump-sum taxes. The budget constraint can therefore be written in terms of tradeables as:

$$F_{t+1} = (1+r) F_t + p_t Y_{N,t} + Y_{T,t} - (p_t C_{N,t} + C_{T,t}) - (p_t I_{N,t} + I_{T,t}) - (p_t G_{N,t} + G_{T,t}) \tag{3}$$

where F is the domestic stock of foreign assets and p the relative price of nontradeables ($p = p_N/p_T$). Equilibrium on the market of nontradeables requires that production equals consumption:

$$Y_N = C_N + I_N + G_N \tag{4}$$

In contrast, consumption of tradeables can be smoothed by international trade. Recursive substitution of (3), using (4), gives the intertemporal budget constraint:

$$\sum\nolimits_{s=t}^{\infty} (1+r)^{t-s} (C_{T,s} + I_{T,s} + G_{T,s}) = (1+r) F_t + \sum\nolimits_{s=t}^{\infty} (1+r)^{t-s} Y_{T,s} \tag{5}$$

[3] Bovenberg and Knaap (2005) is the only study which simulates ageing costs with a detailed CGE model, incorporating a nontradeable sector. Due to adjustment costs of sector-specific capital, the relative price of nontradeables is shown to deviate temporarily from its long-run value during the demographic transition. This outcome is in line with our simulation results in Section 5.

[4] The composition of the sectoral capital stocks does not matter in this analysis.

or, the present value of the expenditures on tradeables must equal the present value of tradeable income. The solutions of two versions of the model are discussed in the next subsections. The first version of the model focuses on the long-run implications of the model. Subsequently, short-run determinants of relative prices are discussed.

2.2 Long-run determinants of relative prices

The Balassa-Samuelson theory assumes perfect international capital mobility and perfect intersectoral factor mobility. These extreme assumptions seem only reasonable over the long run. The classical result that the relative price of nontradeables only depends on supply factors is easily derived, under these conditions.

The assumption of perfect intersectoral labor mobility implies that the wage rate w is equal in both sectors. With a competitive labor market, the marginal productivity of labor in each sector (from (1)) has to equal the common wage rate, or:

$$w = \alpha_N p_N y_N = \alpha_T p_T y_T \tag{6}$$

with $y_x = Y_x / L_x$. Taking logs of (6) yields the long-run equilibrium equation:

$$\ln(p_N/p_T) = \ln(\alpha_T/\alpha_N) + \ln(y_T/y_N). \tag{7}$$

Differentiating (7) results in (with $\hat{X} = d\ln X$)

$$\hat{p} = \hat{p}_N - \hat{p}_T = \hat{y}_T - \hat{y}_N \tag{8}$$

or, the rate of change in the relative price of nontradeables equals the rate of change in the relative average labor productivity (Canzoneri et al., 1999). The estimation of the long-run equation in section 4 is based on (7).

The average productivity of labor y_x might also capture demand shocks. To focus on supply shocks, an expression is derived in terms of the growth rates of *total* factor productivity. From (1) follows that $\hat{y}_x = \hat{\theta}_x + (1-\alpha_x)\,\hat{k}_x$, with $\hat{k}_x = \hat{K}_x - \hat{L}_x$. The marginal productivity condition of capital for a given rental rate implies that $\alpha_x \hat{k}_x = \hat{p}_x + \hat{\theta}_x$. Substitution in (8) gives the classical expression (with $\hat{p}_T = 0$):

$$\hat{p} = \frac{\alpha_N}{\alpha_T}\hat{\theta}_T - \hat{\theta}_N \tag{9}$$

or, the rate of change in the relative price equals the difference between the growth rates of total factor productivity across sectors corrected by the labor shares.

2.3 Short-run determinants

At the other extreme, the short run is represented by assuming perfect factor rigidity. With fixed domestic supply, demand conditions get an effect on the relative price of non-tradeables.

Fixed capital stocks exclude investment ($I_x = 0$). With exogenous production ($\hat{Y}_x = \hat{\theta}_x$) and exogenous government consumption, C_N is fixed by (4). The optimal C_T's follow from maximizing (2), subject to (3). Using (3) to substitute for C_T in (2) yields:

$$U_t = \sum_{s=t}^{\infty} (1+r)^{t-s} \{\phi \ln C_{N,s} + (1-\phi) \ln [(1+r) F_s - F_{s+1} + Y_{T,s} - G_{T,s})]\} \quad (10)$$

From the first order conditions $dU_t/dF_s = 0$, it follows that the optimal C_T is constant, or $C_{T,s} = \bar{C}_T$ ($\forall s$). This constant level can be calculated from (5):

$$\bar{C}_T = \frac{(1+r) F_t + \sum_{s=t}^{\infty} (1+r)^{t-s} (Y_{T,s} - G_{T,s})}{\sum_{s=t}^{\infty} (1+r)^{t-s}} \quad (11)$$

In view of the CD-specification of the instantaneous utility function, expenditures on nontradeables are a fixed fraction of the total budget:

$$p_{N,s} C_{N,s} = \phi \left(p_{N,s} C_{N,s} + p_{T,s} \bar{C}_T \right) \qquad \text{or} \quad (12)$$

$$p_s = \frac{\phi}{1-\phi} \frac{\bar{C}_T}{C_{N,s}} \quad (13)$$

The last equation shows that the level of the relative price depends next to productivity on preferences and government consumption. First, productivity growth in the tradeable and nontradeable sector has an effect via $\bar{C}_T$ and C_N, respectively. With fixed factors, anticipated productivity growth in the tradeable sector has no effect on the *growth rate* of the relative price (since $\hat{C}_T = 0$). Anticipated, as well as unanticipated, productivity shocks in the nontradeable sector are fully transmitted, as $\hat{p} = -\hat{C}_N = -\hat{\theta}_N$ (assuming $\hat{G}_N = \hat{\theta}_N$).

Second, government consumption, assumed to be biased towards nontradeables, obviously affects the relative price. This effect might be better interpreted as a supply effect since it reduces the fraction of output that is available for consumption (see de Ménil, 1994). Noticing that $dG_N = -dC_N$, it is easily verified that $dp/dG_N = -dp/dC_N > 0$.

Third, demographic structure may affect the relative price as well. Older consumers are believed to reveal stronger preferences for nontradeables than younger consumers. The effect of a shift to an ageing population is illustrated by analyzing a shock in the preferences of the representative agent (ϕ). From (11) follows that $d\bar{C}_T/d\phi = 0$ due to the CD-utility specification. The effect on the relative price follows from differentiating (13):

$$\frac{dp_s}{d\phi} = \frac{1}{(1-\phi)^2} \frac{\bar{C}_T}{C_{N,s}} > 0$$

Clearly, following a taste shock in favour of nontradeables, its price must increase when supply is inelastic.[5]

De Gregorio et al. (1994a) analyze the same preference shift in a model with noncompetitive goods and labor markets. The nontradeable sector is assumed monopolistically competitive. Labor is the only production factor and is perfectly mobile across sectors (Financial capital is mobile internationally). The wage is set by a centralized labor union above the competitive level. The labor market is shown to play a crucial role in transmitting shocks to the relative price of nontradeables. An increase in the share of expenditures on nontradeables increases aggregate labor demand. The resulting increase in the wage rate raises the relative price of nontradeables.

Finally, Bergstrand (1991) and De Gregorio et al. (1994b) provide an alternative demand-side explanation for changes in the relative price. By allowing for non-homothetic tastes, the income elasticity of demand is assumed greater (less) than 1 for nontradeables (tradeables). Income growth will therefore cause an increase in the relative price by affecting the demand composition (in the absence of perfect capital mobility). An unattractive feature of their approach is that domestic demand for tradeables is set equal to domestic production; i.e. consumption cannot be smoothed by using the current account.

[5] Rogoff (1992) considers a third case by combining closed capital markets with fixed sector-specific factors. Using $Y_x = C_x + G_x$ ($x = \{N, T\}$) in (12) gives $\hat{p} = \omega_T \hat{\theta}_T - \omega_N \hat{\theta}_N - \left[(\omega_T - 1) \hat{G}_T - (\omega_N - 1) \hat{G}_N \right]$ with $\omega_x = Y_x / C_x$. Next to productivity shocks, demand shocks matter in the determination of the relative price.

3 Data

Sectoral data for 9 main sectors in 16 countries are taken from the STAN-database of OECD (2004b). These sectors are listed in Table 1. The classification into tradeables (T) and nontradeables (N) is based on the ratio of exports to total production (see De Gregorio et al., 1994b). The calculation of these ratios per country is explained in Appendix A. The second column in Table 1 reports the (unweighted) average over all countries in the sample in 2000. Using a threshold of 10% for this ratio, commodities from the sectors agriculture, mining and manufacturing, in addition to transport services are classified as tradeables. The remaining 5 services sectors are considered as nontradeables. The same classification is used by De Gregorio et al. (1994b).

Table 1: The main sectors in OECD (2004b)

abbreviation	E(X/Y)*	classification	description
AG	16.6	T	Agriculture, hunting, forestry and fishing
MI	25.2	T	Mining and quarrying
MA	46.5	T	Total manufacturing
EL	2.0	N	Electricity, gas and water supply
CS	1.7	N	Construction
TD	0.0	N	Wholesale and retail trade; restaurants and hotels
TP	16.9	T	Transport, storage and communication
FB	7.5	N	Finance, insurance, real estate and business services
CM	0.8	N	Community, social and personal services

*unweighted average of the ratio exports/production in 2000 (%).

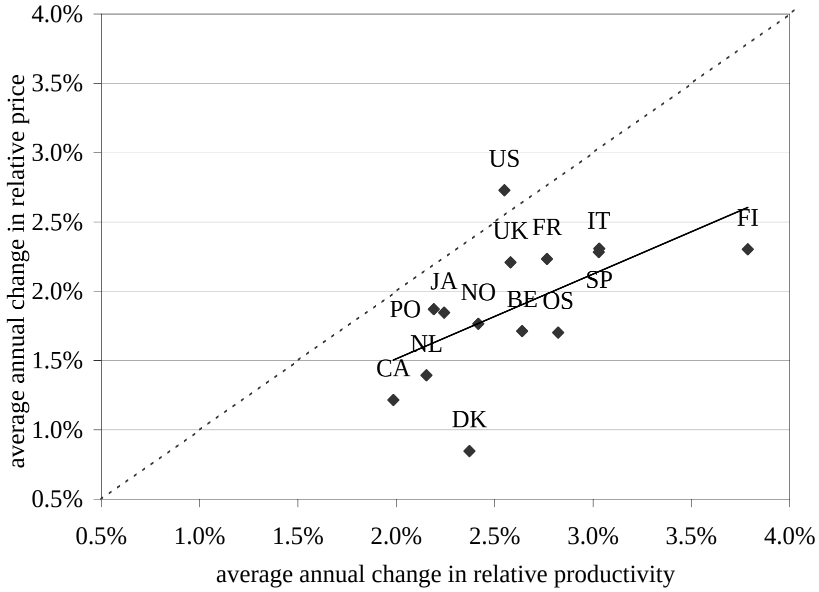
Table 2 lists the countries and the corresponding years that are included in the sample.[6] Differences in data availability over the countries result in an unbalanced panel (In particular, data for Germany are only included after the unification). The Table also shows the substantial share of nontradeable sectors in total value added in 2000. The fraction of nontradeables ranges from 60% in Finland to 75% in the US.

The sources of the basic variables are listed in Table 3. The prices of tradeables and nontradeables are measured by the deflators of value added, or $p_x = va_x/vac_x$ with $x = \{T, N\}$. The relative price of nontradeables is defined as $p = p_N/p_T$.[7] With average

[6] Norway is not included in view of its exceptional share of tradeables (46%), due to the large size of the mining sector (25% in 2000). The estimation of the age effects is sensitive to including this outlier. Including Norway tends to reinforce the ageing effects.

[7] An alternative measure of the relative price is based on sub-indices of consumer prices, see Engel (2002). This approach defines the relative price as the ratio of the consumer price index of commodities and the price index of services. This alternative method suffers from the problem that the classification of

Figure 1: Productivity growth and relative price of nontradeables (1980-2000)



labor productivity computed as $y_x = vac_x/e_x$, relative productivity follows as $y = y_T/y_N$. Canzoneri et al. (1999) provide theoretical and practical justifications for using labor productivity instead of total factor productivity (as in e.g. De Gregorio et al. (1994b)). In particular, the computation of TFP requires in addition data on sectoral capital stocks, which are known to be less reliable (see Froot and Rogoff (1995, note 40)). The fraction of government consumption in GDP is $gs = gc/gdpc$. Data on the age composition of the populations is taken from UN (2001).

A first impression of the long-run relation between p and y is obtained from Figure 1, which plots the average growth rate of relative productivity against the growth rate of relative prices for a common period 1980-2000.[8] The pattern is similar as in De Gregorio et al. (1994b, Fig. 3). The slope of the regression line is positive but different from one, the value predicted by the Balassa-Samuelson hypothesis.

the consumer goods differs from the classification of the production sectors. The correlation between the logs of both measures equals 0.98 and 0.95 for the US and The Netherlands, respectively (see Appendix B).

[8]Except for Japan (1981-2000) and Portugal (1980-1999).

Table 2: Summary statistics per country (%)

	coverage	share in VA of nontradeables in 2000	average growth rates			
			y_T	y_N	y	p
Australia	1974-2001	70.4	3.03	0.92	2.11	1.20
Austria	1976-2002	69.7	3.64	0.98	2.66	1.61
Belgium	1975-2002	72.4	3.74	0.98	2.76	2.11
Canada	1970-2000	65.0	2.44	0.61	1.83	0.43
Denmark	1970-2002	70.2	3.39	0.76	2.63	1.08
Finland	1970-2002	59.6	4.59	1.44	3.16	1.81
France	1978-2002	72.8	3.50	0.86	2.64	2.14
Germany	1991-2002	70.2	3.60	0.55	3.05	1.26
Italy	1970-2002	68.9	3.38	0.61	2.77	1.76
Japan	1981-2001	71.2	3.45	1.36	2.09	1.84
Netherlands	1977-2002	71.3	2.61	0.46	2.15	1.68
Portugal	1977-1999	70.6*	4.29	1.40	2.88	1.71
Spain	1980-2002	69.9	3.02	0.21	2.81	2.20
Sweden	1993-2002	67.5	5.55	0.95	4.59	3.33
United Kingdom	1971-2002	70.8	3.08	0.99	2.09	1.18
United States	1977-2001	75.4	2.87	0.73	2.13	2.40

Source: calculated from OECD (2004b). Notation: y_x is average labor productivity in sector x; $y = y_T/y_N$; p is the relative price $= p_N/p_T$. * in 1999.

Table 3: Source of the basic variables

name	description	source
va_s	value added	OECD (2004b)
vac_s	value added in constant prices	OECD (2004b)
e_s	total employment (persons)	OECD (2004b)
gc	government final consumption expenditure in constant prices	OECD (2004a)
$gdpc$	GDP in constant prices	OECD (2004a)
yc	GDP per head in constant prices and PPPs	OECD (2004a)
pop_a	population shares by age group	UN (2001)

*subscript s denotes sector; a denotes age group (five year bracket).

4 Estimation results

This section presents the estimation results obtained from the panel data discussed in the previous section. Following the theory, we distinguish between long-run equilibrium dynamics (driven by the standard Balassa-Samuelson effect) and the short-run dynamics, incorporating temporary shocks. From an econometric point of view, we run ECM regressions to identify and estimate the (temporary) effects of changes in the population structure on relative prices.

4.1 Non-stationarity and cointegration

Standard theoretical arguments point to an equilibrium relation between relative prices and relative productivity levels. Given that one cannot reject the unit root in either of these series, we use the Engle-Granger procedure to test for a long-run equilibrium relation. Table 4 presents the results for the univariate and the panel cointegration tests. Based on the univariate cointegration tests, we do not find strong evidence in favor of cointegration. Both for the standard Dickey-Fuller (DF) test (not adjusting for serial correlation) and the Augmented Dickey-Fuller (ADF) test we can reject the null hypothesis of no cointegration in 3 out of 16 cases at a significance level of 10%, i.e. for Belgium, Germany and Spain. Restricting the significance level to 5%, we only find cointegration for Germany.

As is well known, however, DF and ADF cointegration tests lack power in small samples. Panel cointegration tests can be used to increase the power of the test. Given that our panel is relatively small, we opt for the standard Im et al. (2003) test. This test has been shown to have reasonable power properties, even in small samples (Pedroni, 1995). Moreover, the test procedure allows for heterogeneity in the cointegrating relationship and can be used in unbalanced panels. The panel version of the ADF test for no cointegration is then given by:

$$W_{\bar{t}_b} = \frac{\sqrt{N}(\bar{t}_b - \bar{\mu}_{\bar{t}_b})}{\sqrt{\bar{\sigma}^2_{\bar{t}_b}}} \tag{14}$$

with $\bar{t}_b = \sum_{i=1}^{N} t_{i,T_i}(p_i)$, where $t_{i,T_i}(p_i)$ denotes the ADF t-test statistic with p_i lags, performed on a sample with T_i observations. The parameters $\bar{\mu}_{\bar{t}_b}$ and $\bar{\sigma}^2_{\bar{t}_b}$ are adjusted to incorporate the unbalanced character of the panel (see Im et al. (2003), remark 3.1) :

$$\bar{\mu}_{\bar{t}_b} = \frac{1}{N}\sum_{i=1}^{N} E(t_{i,T_i}(p_i)), \qquad \bar{\sigma}^2_{\bar{t}_b} = \frac{1}{N}\sum_{i=1}^{N} Var(t_{i,T_i}(p_i)). \tag{15}$$

The population values for $E(t_{i,T}(p_i))$ and $Var(t_{i,T}(p_i))$ have been tabulated by Im et al. (2003). Unlike the standard ADF test statistics, $W_{\bar{t}_b}$ is distributed under the null of no cointegration as a standard normal variate. The last row of Table 4 contains the $W_{\bar{t}_b}$

statistics for our data set. The null hypothesis of no cointegration, i.e. of a unit root in the regression errors, is rejected at the 5% significance level. Based on the panel results we can therefore reject the null hypothesis of no cointegration.

Table 4: Engle-Granger cointegration tests

$$\ln p_{it} = \beta_{i0} + \beta_{i1} \ln y_{it} + \varepsilon_{it}$$

	# obs	ADF[a]	# lags	β_{i0}	β_{i1}	ADF[b] $\beta_{i1}=1$	# lags
Australia	28	−2.312	10	−0.184	0.726	−2.581 *	2
Austria	27	−2.351	4	0.232	0.632	1.634	2
Belgium	28	−3.051 *	3	−0.078	0.712	−0.984	2
Canada	31	−2.774	7	−0.230	0.575	−0.472	4
Denmark	33	−1.797	6	−0.020	0.400	−0.642	2
Finland	33	−2.831	2	−0.123	0.593	0.636	6
France	25	−1.582	2	−0.016	0.813	−0.141	2
Germany	12	−3.787 **	3	0.010	0.363	−4.239 **	3
Italy	33	−2.507	2	0.040	0.647	−3.669 **	2
Japan	21	−1.648	3	0.115	0.885	−2.511	8
Netherlands	26	−2.014	2	−0.212	0.714	−0.578	8
Portugal	23	−1.943	5	0.172	0.646	2.351	9
Spain	23	−3.068 *	5	0.047	0.728	−1.369	9
Sweden	10	−0.705	2	−0.180	0.773	−0.975	2
United Kingdom	32	−1.719	2	−0.308	0.654	−2.030	3
United States	25	−2.119	2	−0.264	1.074	−0.549	2
Panel		−3.540 **				1.421	

[a]tests without trend. Panel test (14) with $\bar{\mu} = -1.386$ and $\bar{\sigma}^2 = 0.985$.

[b]tests without trend. Panel test $\beta_1 = 1$ (14) with $\bar{\mu} = -1.364$ and $\bar{\sigma}^2 = 1.006$.

* and ** denote significance at 10% and 5%, respectively.

Note, however, that we do not recover a long-run one-to-one relation between the (ln) relative price and (ln) relative productivity. Typically, we find a coefficient significantly lower than 1. The exception is the US where we find a coefficient of about 1.07. Moreover, formal tests do reject the homogeneity of this parameter.[9] Although this finding contradicts the Balassa-Samuelson prediction of proportionality, it is not uncommon, e.g. Canzoneri et al. (1999). Table 4 also contains the formal statistics of a cointegration test, assuming a one-to-one relation between relative prices and relative productivity. As

[9]Estimation results are available upon request.

can be observed from the univariate ADF tests, we cannot reject the unit root but in two cases. Performing the panel cointegration test, we find a test statistic of 1.42. Clearly, this test statistic rejects the hypothesis that the cointegration vector satisfies the one-to-one relation between relative prices and productivity.

We conclude that an ECM is the appropriate specification to explain the relative price of nontradeables.[10]

4.2 Modelling the age effects

Next to long-run supply effects, relative prices are affected by temporary shocks as well. We integrate these shocks by means of an ECM model. More in particular, an ECM with fixed effects for the countries is used:

$$\Delta \ln p_{it} = \lambda_i + \beta_1 \ln p_{i,t-1} + \beta_2 \ln y_{i,t-1} + \beta_3 \Delta \ln y_{it} + \beta_4 \Delta gs_{it} + \beta_5 \Delta \ln yc_{it} + \textstyle\sum_{j=1}^{J} \delta_j \Delta pop_{jit} \tag{16}$$

where p_{it} denotes the relative price of nontradeables in country i at time t, λ_i denotes the country-specific fixed effect and y_{it} measures the relative (average) labor productivity differential between tradeables and nontradeables. The short-run effects include, productivity growth differentials $\Delta \ln y_{it}$, changes in government consumption Δgs_{it}, growth rate of per capita income $\Delta \ln yc_{it}$, and demographic changes across J population groups Δpop_{jit}.

In order to model demographic changes we construct $j = 1, ..., J$ population groups. Each population group covers five years and the last cohort considered is 80 years old and over, implying $J = 17$. For each country and each period in time, we compute the fraction of the total population within age group j and denote it by pop_{jit} Given that we concentrate on shocks we incorporate changes in the population fractions in the ECM model.

By decomposing the demographic structure into subgroups covering each five years, we allow for a detailed analysis of changes in the demographics of a country. The drawback of this decomposition is that 17 parameters are needed to model the impact of demographic changes (i.e. δ_j with $j = 1, .., J$). In order to save on degrees of freedom, the approach of Fair and Dominguez (1991) is applied by imposing a quadratic polynomial on the age coefficients:[11]

$$\delta_j = \gamma_0 + \gamma_1 j + \gamma_2 j^2. \tag{17}$$

[10] De Gregorio et al. (1994b) estimate regressions including only variables in first differences. Their results indicate that relative price changes in the short run are mainly driven by demand side factors, whereas relative productivity growth differentials become the dominant determinant in the long run. These dynamics are better captured in an ECM.

[11] The approach is explained for a quadratic polynomial. The definitions are easily adapted for a polynomial of a different order.

The main advantage is that instead of estimating seventeen parameters we only estimate three, γ_0, γ_1 and γ_2. The implied subgroup-specific sensitivities, δ_j, can then be recovered by (17) for each specific subgroup j. Note that by construction, changes in population shares add up to zero and hence, the γ_0 parameter is not identified in our setting. We use this additional degree of freedom to normalize the age impacts to have mean zero: $\frac{1}{J}\sum_j \delta_j = 0$ implying a value for γ_0 of:

$$\gamma_0 = -\frac{1}{J}\left(\gamma_1 \textstyle\sum_j j + \gamma_2 \sum_j j^2\right). \tag{18}$$

Using the polynomial expansion for the parameters (equation (17) in the ECM (equation (16)) we obtain the estimated ECM:

$$\Delta \ln p_{it} = \lambda_i + \beta_1 \ln p_{i,t-1} + \beta_2 \ln y_{i,t-1} + \beta_3 \Delta \ln y_{it} + \beta_4 \Delta gs_{it} + \beta_5 \Delta \ln yc_{it} + \textstyle\sum_{k=1}^{2} \gamma_k \Delta f_{kit} \tag{19}$$

with the auxiliary variables Δf_{kit} defined by:

$$\Delta f_k = \textstyle\sum_j^J j^k \Delta pop_j - \frac{1}{J}\sum_j^J j^k \qquad \text{with } k = 1, 2. \tag{20}$$

Although neither of the auxiliary variables has a proper economic interpretation, they are instrumental in estimating the parameters γ_1 and γ_2, which in their turn generate the group sensitivities δ_j $(j = 1, ..., J)$.

Although the Fair-Dominguez approach allows for the estimation of detailed age effects, it does not solve the multicollinearity problem between the age variables, see Lindh and Malmberg (2000).[12] As a result, the approach might suffer from imprecise estimates. As an alternative, we also include more aggregated demographic measures, namely the standard dependency ratios. The old-age dependency ratio is defined as the ratio of the number of persons aged 65 and over to the number of persons of working age (20-64). Similarly, the young-age dependency ratio is the ratio of the number of young persons (under 20) to the number of persons of working age:

$$dep_y = \frac{\sum_{j=1}^{4} pop_j}{\sum_{j=5}^{13} pop_j} \qquad , \qquad dep_o = \frac{\sum_{j=14}^{17} pop_j}{\sum_{j=5}^{13} pop_j}.$$

In the following section we compare the results obtained following the two approaches for modelling the age effects.

[12] The correlation between Δf_1 and Δf_2 equals 0.998 in our sample.

4.3 Results

Table 5 shows the results for two panel sets of countries and three specifications of the age effects. The full sample contains all 16 countries, while the second sample only consists of 12 European countries, excluding the UK. The latter sample consists of economies which are believed to have less flexible markets, leading to a slower adjustment of the relative price of nontradeables. Three specifications for the age effects are estimated. The first two specifications use the Fair and Dominguez (1991) approach with respectively a linear ($\gamma_2 = 0$ in eq. (17)) and a quadratic polynomial (eq. (17)). To check the robustness of estimation results, a third version is estimated with the Fair and Dominguez population variables replaced by two dependency ratios, defined above.

First, we discuss the estimates of the non-age coefficients in the full sample. Table 5 indicates that these estimates are robust to the specification of the age effects. The coefficient of $\ln p_{-1}$ shows that the relative price adjusts relatively slowly to its long-run level. A parameter value for β_1 of -0.15 implies a halving time of about $\ln(0.5)/\ln(0.85) =$ 4.3 years. Relative productivity has a strong, positive long-run effect on the relative price of nontradeables. In the full sample, the implied long-run coefficient $(-\beta_2/\beta_1)$ equals 0.87, 0.79 and 0.76 for the respective age effect specifications (the standard errors are around 0.06).

The estimated short-run effects correspond to previous findings in the literature. The change in (ln) relative productivity has a significant, positive impact in the same year. A growth rate of relative productivity of 1% increases the relative price by 0.26% in the short run. Its lagged growth rate is significant in none of the cases. As government expenditures are mainly spent on nontradeables, a positive effect on the relative price is expected. Increasing the share of government consumption in GDP by 1% is found to raise the relative price by 1.7% in the current year and by 0.8% in the next year. This is comparable to the total effect reported in De Gregorio et al. (1994b), ranging from 1.5 to 2%.[13] The growth rate of income per capita is only significant when lagged one year in the first two cases.[14] De Gregorio et al. (1994b) report a positive effect (around 0.3) of the contemporaneous income growth rate, while the evidence on this variable is mixed in Lane and Milesi-Ferretti (2002).[15]

Next, we focus on the age coefficients in the full sample.[16] We find evidence in favour of

[13] De Gregorio et al. (1994a) found mixed results for this variable for the five large EU-countries, probably due to the small samples.

[14] The income variable is measured in PPPs. Similar results are obtained when yc is measured in constant exchange rates.

[15] Extending the specification with more than one-year lags of the three short-run variables was rejected

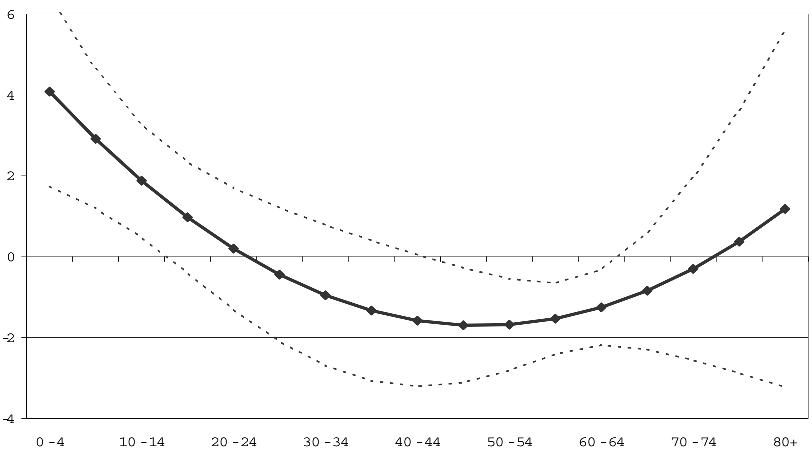
[16] Lagged age variables are not included in view of the multicollinearity problem

Table 5: Estimation results*

	Linear age effects		Quadratic age effects		Dependency ratios	
Full Sample (#obs=378)						
$\ln p(-1)$	−0.1524	(0.0267)	−0.1554	(0.0266)	−0.1735	(0.0285)
$\ln y(-1)$	0.1328	(0.0195)	0.1235	(0.0200)	0.1327	(0.0197)
$\Delta \ln y$	0.2604	(0.0564)	0.2502	(0.0564)	0.2649	(0.0563)
Δgs	1.7219	(0.3710)	1.7419	(0.3697)	1.6526	(0.3697)
$\Delta \ln yc$	0.1703	(0.0981)	0.1879	(0.0982)	0.1430	(0.0975)
Δf_1	−0.3373	(0.1165)	−1.3657	(0.5420)		
Δf_2			0.0658	(0.0339)		
Δdep_y					1.0074	(0.3356)
Δdep_o					−0.2935	(0.5650)
$\Delta \ln y(-1)$	−0.0970	(0.0519)	−0.0965	(0.0517)	−0.0886	(0.0517)
$\Delta gs(-1)$	0.8130	(0.3729)	0.8534	(0.3721)	0.7770	(0.3717)
$\Delta \ln yc(-1)$	0.1903	(0.0961)	0.2075	(0.0961)	0.1727	(0.0957)
European Continent (#obs=262)						
$\ln p(-1)$	−0.1423	(0.0340)	−0.1493	(0.0349)	−0.1426	(0.0363)
$\ln y(-1)$	0.1102	(0.0224)	0.1095	(0.0225)	0.1037	(0.0228)
$\Delta \ln y$	0.2483	(0.0575)	0.2489	(0.0575)	0.2455	(0.0582)
Δgs	2.2887	(0.3697)	2.3082	(0.3704)	2.1576	(0.3669)
$\Delta \ln yc$	0.3139	(0.0917)	0.3200	(0.0919)	0.2812	(0.0922)
Δf_1	−0.2320	(0.1049)	−0.7502	(0.5661)		
Δf_2			0.0331	(0.0356)		
Δdep_y					0.4496	(0.3446)
Δdep_o					−0.1664	(0.5467)

*Fixed effects for countries are included. Standard errors in brackets.

Figure 2: Age effects with quadratic specification in full sample (with 5% confidence interval)



demographic effects on relative prices.[17] The age coefficients in the linear and quadratic specifications are statistically significant (The p-value of the Δf_2-coefficient equals 5.3%). The negative coefficient of the linear specification already indicates that the relative price increases with the share of younger generations. The implications of the quadratic specification are better understood after converting the estimates by (17) into the age coefficients (δ), given in Figure 2. The two standard error bounds are calculated per single coefficient.[18] The age effects, implied by the quadratic form, should be interpreted as follows: an increase in the share of age class i by 1% changes the relative price by δ_i%. Although we find evidence of a significant effect of demography on relative prices, it is (surprisingly) not the older but the younger generations that affect the relative price. The first three age classes (aged under 15 years) have a significant upwards effect on the price, whereas elderly (older than 65 years) have a insignificant effect.

The same pattern is found in the last case with dependency ratios. The change in the young-age dependency ratio has a significant positive impact on the relative price, whereas an insignificant effect is found for the change in the old-age ratio.[19] The relevance of the age composition is further illustrated by way of simulations in the next section.

The findings are altered when the sample is limited to the 12 European countries (except

[17] One could argue that government expenditures, like on public education and health services, have an important age component. However, dropping the variable Δgs hardly affects the estimated age profiles.

[18] Therefore, the confidence interval might be somewhat misleading since the bounds cannot hold simultaneously for all coefficients in view of the zero-sum restriction.

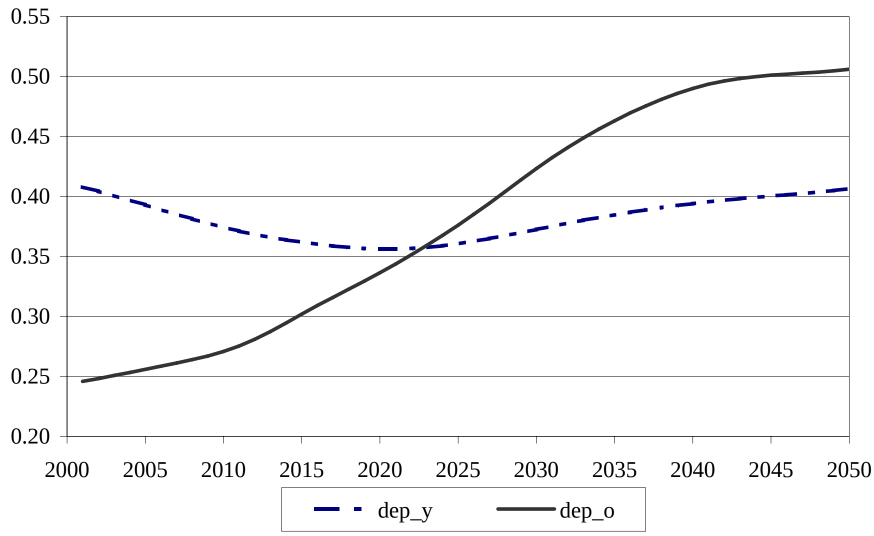
[19] An experiment with country-specific coefficients for Δdep_y and Δdep_o shows that 14 and 9 (out of 16) of these are positive, respectively.

the UK). First, the long-run effect of relative productivity growth is smaller, ranging from 0.73 to 0.77 (with std. err. around 0.06). Second, lags for the three short-run variables are rejected. Third, the contemporaneous growth rate of income per capita has a significant, positive effect in the three cases. Finally, the age coefficients are no longer significant, except for the linear case.

5 Simulations

The estimation results provide some evidence that the relative price of nontradeables is affected by the demographic composition. In view of the substantial demographic shifts projected for the coming decades, the age structure might become a more important determinant of the short-run adjustments of the relative price. Figure 3 shows the dependency ratios, calculated for the projected population living in all countries in the full sample. Projections are for the central variant in UN (2001). As is well-known, the old-age dependency ratio is projected to double over the period 2000-2050. The projected young-age ratio remains relatively stable. We use these projections to simulate the effects of ageing on relative prices. Obviously, projecting the relative price 50 years ahead is a perilous exercise. The simulations only aim to indicate the potential size of the contribution that ageing might have to the development of the relative price.

Figure 3: Projected dependency ratios in full sample



Simulations are performed for the full-sample specification using the dependencies ratios, reported in Table 5. Given the statistical insignificance of the coefficient of dep_o, this parameter is set to zero.[20] The future values of the exogenous variables are generated as follows:

- The relative price in the base year (2000) is normalized at one ($p_0 = 1$).
- Relative productivity is calculated by imposing long-run equilibrium, or $\ln y_0 = -\lambda_1/\beta_2$. The future series is constructed by fixing the growth rate of output per worker at the sample mean, or $\Delta \ln y = 2.55\%$ ($t > 0$).
- The conservative assumption is made that the share of government consumption is not affected by ageing ($\Delta gs = 0$).
- Income per capita is assumed to grow at the sample mean of $\Delta \ln yc = 2.11\%$.
- Changes in the dependency ratio (Δdep_y) are based on the UN-population projections summed over the OECD-subset of 16 countries.

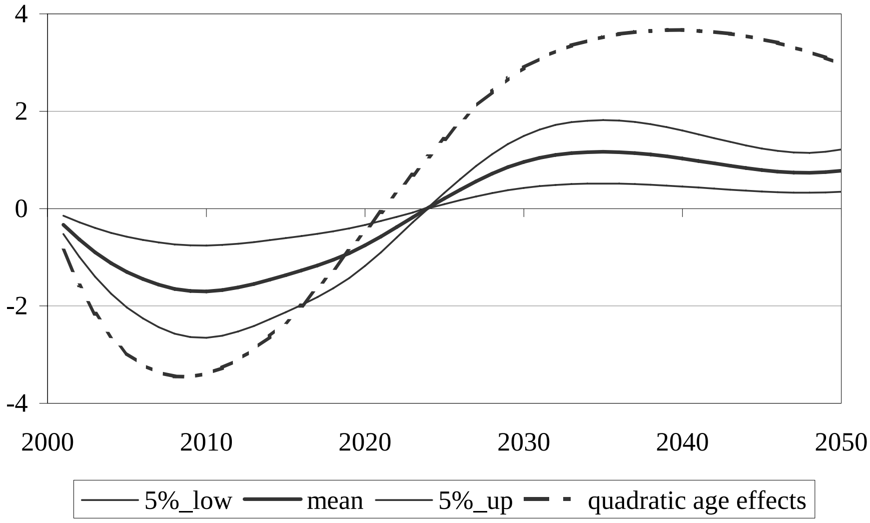
To isolate the age effects, dynamic simulations for p are performed without (i.e. $\Delta dep_y = 0$) and with changes in the dependency ratio. The confidence interval is calculated by running a Monte Carlo experiment on the age coefficient. Based on the estimated coefficient and its variance, alternative values for the age coefficient are drawn for 2000 replications.[21] After the whole path is generated for each replication, results are ordered per year. The 10% confidence interval is represented by taking the lowest 5% and the highest 5% of the outcomes.

Figure 4 reports the percentage difference between the result with and without age effects. The path of the relative price reflects the small changes in the young-age ratio. The deviation between both scenarios reaches its maximum in 2035 (= 1.17%), implying that the age effects raise the average annual growth rate over the period 2000-2035 by only 0.06% (points). Obviously, the scenarios converge once the demographic structure stabilizes. A similar pattern is obtained when the model with quadratic age effects is stimulated, although the effects (and confidence bounds) are much larger.

[20] Re-estimation gives similar coefficients as reported in Table 5 (e.g. the coefficient of Δdep_y is 0.98). Simulation with the latter model therefore yields similar outcomes but with a larger confidence interval. The constant term of the US is used ($\lambda_1 = -0.03$). Since we are primarily interested in the difference between scenarios, the choice of this value does not matter.

[21] Notice we neglect the uncertainties arising from the other coefficients and from the population projections.

Figure 4: Simulated age effects on relative price of nontradeables; case with young-age dependency ratio in full sample (%)



6 Long-run effects: demographics and relative productivity

The analysis sofar assumed that demographic shifts only affected the relative price in the short run. This section extends the impact of demographic factors to the long run by estimating the relation between relative productivity ($y = y_T/y_N$) and the age structure of the population. Cutler et al. (1990) find that a decrease in the labor force growth raises the growth of aggregate labor productivity. More sector-specific results are not available in the literature. We have estimated a relation between the level of relative productivity and both dependency ratios, next to a trend term.

$$\ln y_{it} = \omega_{i0} + \omega_1 trend_t + \omega_2 dep_{y,it} + \omega_3 dep_{o,it}. \tag{21}$$

Appendix C shows that the panel test rejects the hypothesis of no cointegration when the dependency ratios are included in (21). Table 6 reports the estimation results of the cointegrating relation between demographic factors and relative productivity. Table 6 shows that labor productivity in the tradeable sector is estimated to grow exogenously 2.7% faster than productivity in the nontradeable sector. Also, both demographic variables have a positive, significant effect on relative productivity in both samples (except dep_y in the small sample). The results indicate that in an ageing population with a growing share of inactive age groups, labor productivity in the tradeable sector increases relatively to productivity in the (labor-intensive) nontradeable sector. In both samples, the old-age dependency ratio has a larger effect than the young one. When demographic changes are believed to have a long lasting effect, ageing might have a much more profound effect on projections of the relative price.

Table 6: Estimation results level equation for (ln) relative productivity*

	Full sample		European Continent	
trend	0.0266	(0.0008)	0.0273	(0.0008)
dep_y	0.3180	(0.0899)	0.1686	(0.0972)
dep_o	0.6080	(0.1936)	0.9724	(0.2049)
#obs	410		273	

*Fixed effects for countries are included.

Standard errors in brackets.

Based on this long-run relation, the ECM (eq. (16)) is re-estimated with the fitted values $\ln \widehat{y}$ of (21), while observations are still used for the growth rates $\Delta \ln y$.[22] The dependency ratios are only included in the long-run equation. The estimates in Table 7 are similar to Table 5.

Table 7: Estimation results ECM* ($\ln y$ dependent on dependency ratios)

	Full Sample		European Continent	
$\ln p_{-1}$	-0.1017	(0.0207)	-0.1013	(0.0273)
$\ln \hat{y}_{-1}$	0.0909	(0.0151)	0.0829	(0.0181)
$\Delta \ln y$	0.2105	(0.0516)	0.1936	(0.0556)
Δgs	1.8322	(0.3432)	2.0606	(0.3634)
$\Delta \ln yc$	0.2090	(0.0868)	0.2797	(0.0903)
#obs	394		262	

*Fixed effects for countries are included.

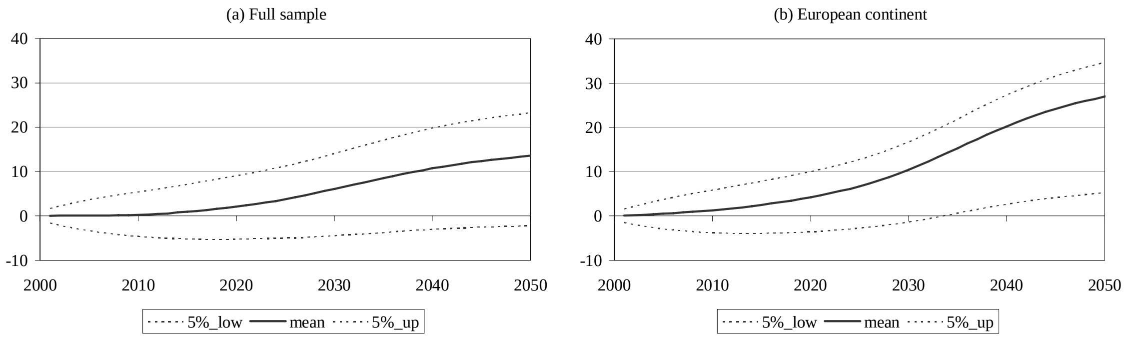
Standard errors in brackets.

The estimates reported in Table 7 are now used to perform simulations, following the methodology discussed in Section 5. The starting level of $\ln y_0$, and thus $\ln p_0$, is equalized between the cases without ($\omega_2 = \omega_3 = 0$) and with age effects by re-scaling the constant term of (21) in the case without age effects. The growth rate of productivity $\Delta \ln y$ is calculated from the simulated $\ln y$. Figure 5 shows that the rising old-age dependency ratio now dominates the deviation between the scenarios, resulting in an extra increase of the relative price by 15.1% and 31.2% in the long run for the full and small samples, respectively. The larger effect for the European countries is explained by the larger coefficient of

[22] In another variant, fitted values are used for the level *and* the change of $\ln y$. Estimation results are reported in Table C.3.

dep_o in Table 6, combined with a stronger rise in the projected old-age dependency ratio (it more than doubles from 27% in 2001 to 61% in 2050).

Figure 5: Simulated demographic effects on relative price of nontradeables (%) ($\ln y$ dependent on dependency ratios)



Ageing contributes to an increase of the average annual growth rate of the relative price by 0.32% and 0.57% over the period 2000-2035. This price increase feeds general inflation. Under the assumption that the consumer price index is a geometric average of the price of tradeables and nontradeables, the inflation rate is calculated as $\hat{P} - \hat{p}_T = \omega_N \left(\hat{p}_N - \hat{p}_T\right)$, where ω_N represents the share of nontradeables in consumption. An increase in the growth rate of the relative price by 0.32% (points), combined with a consumption share of 70%, implies an increase in inflation by 0.4% per annum.[23] However, as indicated by the confidence bounds, this upward effect is still highly uncertain.

[23] Lindh and Malmberg (2000) report for the OECD a significant relation between general inflation and (log) age shares. The results are consistent with the expectation that net savers weaken and net consumers strenghen inflationary pressures.

7 Concluding remarks

Estimations on a panel of OECD-countries provide evidence that the relative price of nontradeables depends on the age structure of the population, even after correcting for the standard explanatory variables.[24] Statistical tests support the theoretical insight that the long-run level is exclusively determined by relative productivity growth differentials. We first focus on the case in which age shares can only have a transitory effect on the relative price. The relative price is found to increase with the share of younger generations, while older generations (aged over 65 years) have an insignificant effect. However, a simulation with demographic projections shows that the impact of ageing probably remains small. In contrast, the second case allows for permanent age effects by considering a link between the level of relative productivity and the age composition. An increase in both, young and old-age, dependency ratios is found to raise, via relative productivity, the relative price. In this case, the demographic transition in the coming decades might have substantial and long-lasting consequences for the relative price.

When age effects are believed to matter in the determination of the relative price, three implications of ageing are identified. First, ageing might contribute to inflationary pressures. As demographic changes are relatively easily predictable, inflation forecasts might be improved by incorporating this information (see Lindh and Malmberg, 2000). Second, elderly are in particular harmed by the increase in the relative price since they spend a larger fraction of their budget on nontradeables (Hobijn and Lagakos, 2003). Higher prices might also aggravate the sustainability problems of pension systems that adjust benefits for inflation. Although these implications are theoretically sound, our empirical results suggest that the inflationary pressure due to ageing remains very uncertain.

Third, the real exchange rate, defined as the ratio of the aggregate price levels, is a function of the relative price of nontradeables in the two countries. A country in which ageing proceeds faster than in its trading partners might be confronted with an appreciation of the bilateral real exchange rate.[25] However, evidence suggests that real exchange rates are hardly related to relative prices of nontradeables, see e.g. Engel (1999).[26] Therefore, the age-specific effects found for the relative price of nontradeables are unlikely to affect real exchange rates.

[24] Our study might suffer from the same problem that Poterba (2001) faced in the estimation of the relationship between the demographic structure and asset returns. It might be difficult to find strong evidence of age-specific effects simply because of a lack of observations on ageing populations.

[25] The estimates in Andersson and Osterholm (2001) support the hypothesis that age shares have a significant effect on the Swedish (log) real exchange rate. Young adults (15-29 years) and retirees (65+) seem to have an appreciating effect, while prime and middle aged (30-64) have a depreciating effect.

[26] In contrast, Obstfeld and Rogoff (2004) argue that the requisite depreciation of the real dollar exchange rate will be dominated by changes in the relative price of nontradeables due to the very large production share of the latter sector.

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Appendix A

Data on production and exports of goods in current prices are taken from OECD (2004b). For the calculation of the export shares of the services sectors, exports data from OECD (2003) are combined with production data from OECD (2004b). The mapping between the different sector classifications is defined in Table A.1. Table A.2 gives the resulting export shares for 2000. Notice that exports can exceed production for several reasons: (i) exports may include re-exports; (ii) production data are activity-based, while trade statistics are product-based and (iii) production and exports may be valued in different prices (see OECD, 2004b).

Table A.1: Mapping for service sectors in OECD (2003)

classification in OECD (2003)	mapping
Transportation	TP
Travel	-
Communications services	TP
Construction services	CS
Insurance services	FB
Financial services	FB
Computer and information services	FB
Royalties and license fees	-
Other business services	FB
Personal, cultural and recreational services	CM
Government services	CM

Table A.2: Ratios Exports/Production in 2000 (%)

	AG	MI	MA	EL	CS	TP	FB	CM
Australia*	-	-	20.8	-	-	-	-	-
Austria	6.1	13.1	63.3	4.5	2.2	18.7	11.5	0.9
Belgium	54.9	-	98.1	3.1	2.8	30.1	23.0	2.6
Canada	22.8	50.5	52.7	11.2	0.3	11.2	5.8	1.1
Denmark	25.8	56.7	66.2	2.2	0.0	46.6	6.2	0.0
Finland	4.8	14.1	48.9	0.2	2.8	10.3	4.4	0.2
France	14.5	-	37.8	5.8	2.0	13.1	4.1	0.6
Germany	11.9	9.3	45.0	0.8	2.0	10.1	4.6	0.7
Italy	9.0	6.8	34.1	0.0	1.2	7.2	3.9	0.4
Japan	0.4	1.2	16.5	0.0	0.8	5.1	1.3	0.1
Netherlands	50.6	44.7	84.1	0.5	4.7	41.3	12.5	1.3
Portugal*	3.9	-	37.8	0.6	-	-	-	-
Spain*	19.4	12.3	29.3	0.3	-	-	-	-
Sweden	8.0	46.8	52.5	0.5	3.9	12.3	9.4	0.4
United Kingdom	7.2	44.9	40.3	0.0	0.0	9.4	9.7	0.6
United States	9.9	2.5	16.7	0.1	0.0	4.5	1.5	0.9
Average (unweighted)	16.6	25.2	46.5	2.0	1.7	16.9	7.5	0.8

Source: OECD (2003, 2004b) (exports are zero for sector TD). *Figures for 1999.

Appendix B

Engel (2002) applies different measures of the relative price of nontradeables. This appendix compares the relative value added deflator with the relative consumer price. In the latter approach, the price of tradeables is calculated as the geometric weighted average of the price index of food and the price index of all goods excluding food. The price of nontradeables is measured as the average of the price index of services less rent and the price index of rent. The alternative relative price is defined as:

$$p_{alt} = \frac{(p_f)^{\frac{\beta_f}{\beta_f+\beta_g}} (p_g)^{\frac{\beta_g}{\beta_f+\beta_g}}}{(p_s)^{\frac{\beta_s}{\beta_s+\beta_r}} (p_r)^{\frac{\beta_r}{\beta_s+\beta_r}}}$$

with $\sum_j \beta_j = 1$. Data are taken from OECD, Main Economic Indicators. For the US, the weights β_i in 2001 are found in Engel (2002, note 9): 0.16, 0.26, 0.27 and 0.31. Weights for The Netherlands are fixed at the 1985 value: 0.15, 0.44, 0.19 and 0.22. The correlation between the logs of both measures equals 0.98 and 0.95 for the US and The Netherlands, respectively.

Figure B.1: The relative price of nontradeables in the US (in log)

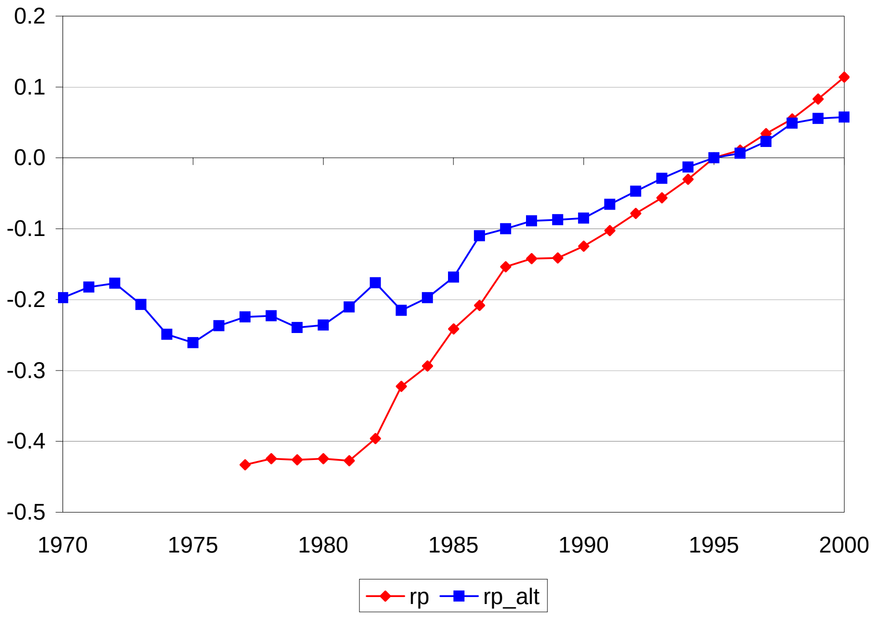
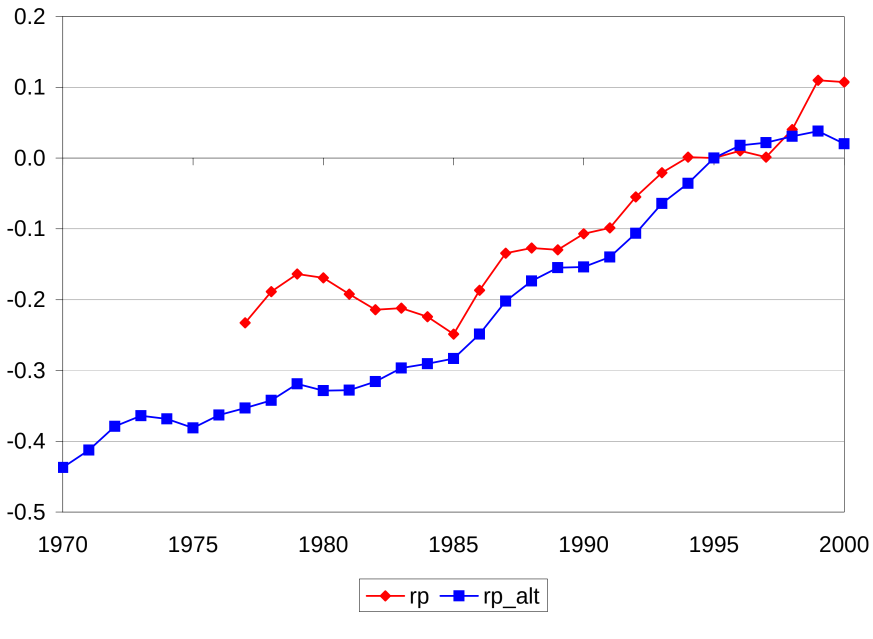


Figure B.2: The relative price of nontradeables in The Netherlands (in log)



Appendix C

Table C.1: Engle-Granger cointegration tests

$$\ln y_{it} = \omega_{i0} + \omega_{i1} trend_t + \omega_{i2} dep_{y,it} + \omega_{i3} dep_{o,it} + \varepsilon_{it}$$

	# obs	ADF (no age)[a]	# lags	ADF (with age)[b]	# lags
Australia	28	-3.27585^{*}	10	-3.41439	2
Austria	27	1.72504	10	-1.52137	10
Belgium	28	-2.86578	8	-2.49329	8
Canada	31	-1.88946	3	-2.59904	3
Denmark	33	-1.99517	10	-2.01343	7
Finland	33	-1.28112	2	-3.49070	2
France	25	-2.28915	9	1.80307	9
Germany	12	-1.35875	3	-4.72813^{**}	3
Italy	33	-0.69351	2	-2.32331	2
Japan	21	-2.16127	3	-1.38994	5
Netherlands	26	-3.91448^{**}	10	-1.92532	10
Portugal	23	-2.25554	8	-1.85048	8
Spain	23	-1.46022	2	-2.30047	3
Sweden	10	-1.03784	2	-2.35567	2
United Kingdom	32	-2.54010	3	-2.79104	3
United States	25	-1.12571	5	-1.81422	9
Panel		-1.66907^{*}		-3.27870^{**}	

[a] Panel test (14) with $\bar{\mu} = -1.336$ and $\bar{\sigma}^2 = 1.118$.

[b] Panel test (14) with $\bar{\mu} = -1.334$ and $\bar{\sigma}^2 = 1.116$.

* and ** denote significance at 10% and 5%, respectively.

Table C.2: Coefficients of LR-relation

$\ln y_{it} = \omega_{i0} + \omega_{i1} trend_t + \omega_{i2} dep_{y,it} + \omega_{i3} dep_{o,it} + \varepsilon_{it}$

	no age effects		with age effects			
	ω_0	ω_1	ω_0	ω_1	ω_2	ω_3
Australia	−0.39221	0.02124	−2.58865	0.01124	1.04934	9.87727
Austria	−1.11853	0.02703	−1.80816	0.04292	2.20277	−2.28969
Belgium	−0.64596	0.02576	0.37739	0.02103	−1.27291	−1.47452
Canada	−0.08486	0.01657	−0.1991	0.03474	0.93898	−3.88974
Denmark	−0.65553	0.02394	−1.99941	0.03465	1.85038	1.34163
Finland	−0.79111	0.03176	−2.05853	0.04399	2.31984	0.00090
France	−0.69913	0.02745	−1.78705	0.03101	1.15526	1.85583
Germany	−0.88732	0.03019	−4.05177	0.03322	9.04696	0.02830
Italy	−0.89238	0.03061	−0.38738	0.02523	−0.68636	−0.45011
Japan	−0.63515	0.01977	−1.37451	0.02529	1.09198	0.79560
Netherlands	−0.25283	0.02125	1.66202	0.04726	1.46021	−14.90500
Portugal	−0.94745	0.02682	−2.32901	0.06936	2.35292	−3.02525
Spain	−0.89482	0.03178	−3.85115	−0.02671	0.47877	16.82919
Sweden	−1.18503	0.04505	3.87917	0.03124	0.01084	−15.61256
United Kingdom	−0.30972	0.02181	1.35384	0.01634	−1.41772	−3.38093
United States	−0.37615	0.02368	−0.76996	0.01993	−0.32567	3.09473

Table C.3: Estimation results ECM* ($\ln y$ dependent on dependency ratios)

	Full Sample		European Continent	
$\ln p_{-1}$	−0.1198	(0.0229)	−0.0951	(0.0292)
$\ln \hat{y}_{-1}$	0.1074	(0.0165)	0.0776	(0.0188)
$\Delta \ln \hat{y}$	3.7495	(1.8669)	0.0150	(0.5165)
Δgs	1.3329	(0.3697)	1.7253	(0.3599)
$\Delta \ln yc$	0.1920	(0.0981)	0.2802	(0.0936)
$\Delta \ln \hat{y}_{-1}$	−3.1855	(1.8249)		
Δgs_{-1}	0.8337	(0.3646)		
$\Delta \ln yc_{-1}$	0.1141	(0.0964)		
$\#obs$	378		262	

*Fixed effects for countries are included. Standard errors in brackets.
Fitted values are used for $\ln \hat{y}$ and $\Delta \ln \hat{y}$.