Factor prices and endowments in a globalised world*

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Abstract

This paper explores empirically the relationship between relative wages and relative endowments of skilled and unskilled workers in open economies across the world. It is guided by a version of Heckscher-Ohlin (HO) theory that has emerged from recent analytical work on trade and is more general (so labelled GHO) than the standard HOS model. In either HO model, openness to trade should benefit a country's more abundant factor, a prediction for which we find support both across and within countries. A weakness of the HOS model, however, is its inability to explain the observed sensitivity of wages to variation in endowments in open economies, especially within countries over time. This sensitivity is consistent with the GHO model, in which, as shown by our results, the impact of endowments on wages depends on the height of barriers to trade and on the share of wages in the cost of production. In brief, greater openness to trade seems to reduce - rather than to eliminate - the effect of variation in endowments on factor prices.

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I Introduction

A valuable but largely unrecognised by-product of recent advances in the economic analysis of trade is a specification of Heckscher-Ohlin (HO) theory that is more general than the usual Heckscher-Ohlin-Samuelson (HOS) model, which it nests as a special case. This more general model, which we label GHO, was implicit in the writings of Ohlin himself and in some earlier applications of HO theory to explain the commodity composition of trade and the effects of trade on factor prices. In this paper, we use the GHO model to guide an empirical investigation of worldwide relationships between factor prices, endowments and trade openness.

The distinguishing characteristic of GHO, as compared to HOS, is that from the perspective of an individual country, demand and supply in world markets are less than infinitely elastic. A longstanding explanation of inelasticity, due to Armington (1969) and embodied in many CGE and gravity models (Anderson, 1979), and in Krugman (1979), is qualitative differences among the varieties of goods made by different countries or firms. More recently, however, inelasticity has emerged as a feature of models with stochastic variations in efficiency among countries (Eaton and Kortum, 2002) and firms (Melitz, 2003). As shown by Arkolakis et al. (2012) and Costinot and Rodrìguez-Clare (2014), these alternative sources of inelasticity are interchangeable for normative analysis of welfare gains from reduction of trade costs. This interchangeability applies also to positive analysis, and is the basis of the GHO model (which collapses to HOS if elasticities are assumed to be infinite).

The purposes of this paper are to set out a basic GHO model, extended to include non-iceberg trade (and trade-related) costs, that explains the determination of factor prices in open economies, and to compare its predictions with regularities observed in the data. For lack of data on the prices of other immobile factors, we focus (like many other studies) on the wages of skilled and unskilled workers, using panel data from the recently developed World Input-Output Database (WIOD), which covers 40 countries during 1995-2009.

Our empirical results provide support for the principles of HO theory, and in particular for the prediction that lower barriers to trade tend to raise the earnings of a country's abundant factors relative to those of its scarce factors, both across countries (a dimension that few earlier studies have been able to examine) and over time within countries (where our results are more clear-cut than those of most other recent studies). However, we also find a strong inverse relationship between the relative wages and relative endowments of skilled and unskilled workers, which is not consistent with the canonical HOS model. Multiple cones of diversification could reconcile an inverse cross-country wage-endowment relationship with the HOS model, and we find some - albeit weak - evidence for their existence. What flatly contradicts the HOS model is the sensitivity of relative wages to changes in relative skill supplies within countries over time, a pattern found also in other studies.

This evidence of factor price sensitivity, however, is consistent with the GHO model, in which relative wages are predicted to vary with relative skill supplies, even in an economy that is open to trade, because of the inelasticities of demand and supply mentioned above. The degree of sensitivity depends on the magnitude of what Arkolakis et al. (2012) call 'trade elasticities' - the responsiveness of imports in individual sectors to changes in their purchaser prices relative to those of home-produced goods. Our empirical results also suggest that the sensitivity of relative wages to relative skill supplies is smaller in countries and periods in which there is greater openness to trade and in which non-iceberg trade and trade-related costs are smaller - patterns that the GHO model can rationalise. If home suppliers have smaller shares of the home market, as a result of low trade barriers, the outcome is less affected by elasticities of substitution among goods (which are lower than among varieties within sectors). In addition, thinner non-iceberg trade cost wedges increase the responsiveness of purchaser prices to wages, which enables factor-market clearing adjustments of output mix to be achieved with smaller changes in wages.

At a theoretical level, our paper extends the multi-factor multi-sector analysis of Costinot and Rodrìguez-Clare (2014, p. 221-3) and related work of Burstein and Vogel (2011) by including non-proportional variable trade (and trade-related) costs. This extension is based on Wood (2012), which links back to Alchian and Allen (1964) and Hummels and Skiba (2004), and in the present paper is widened to include discussion of the effects of traded intermediates and internationally mobile capital (as in Wood, 1994).

Empirically, we are the first explicitly to apply the GHO model to data on the determination of factor prices. Our paper thus adds to Romalis (2004) and Chor (2010), who both used an essentially similar model (without the GHO label) to analyse the effect of factor endowments on the composition of trade. Though relying on different reasons for trade inelasticity, their common contribution was to support the HO explanation of trade patterns while avoiding the indeterminacy or extreme specialisation of HOS models with more goods than factors.[1]

Our paper also adds to empirical studies of non-equalisation of factor prices associated with multiple cones of diversification (e.g. Davis and Weinstein, 2001; Schott, 2003; Debaere

[1]This escape is possible because in GHO differences in factor intensity among goods affect the composition of output not only through their role in clearing factor markets (as in HOS) but also by influencing the relative prices of goods which, with inelastic demand, influence their relative sales.

and Demiroglu, 2003; Xiang, 2007; Kiyota, 2011, 2012a). Most of these studies, however, have been limited to the quantity side (relating differences among economies in product mix or choice of technique to differences in their endowments), with only Kiyota also analysing factor price differences and then only within one country. Our paper is the first, we believe, to test for multiple cones in cross-country factor price data.

Most studies of factor price equalisation - always rejected by their results - have also been within single countries (Hanson and Slaughter, 2002; Tomiura, 2005; Bernard et al., 2008, 2013). Data on wages in many countries are used by Blum (2010) to reject the HOS factor price insensitivity theorem - the relative wages of skilled workers fall over time when relative skill supplies rise - and by Marshall (2012). We extend their work by explicitly considering the effect of trade barriers, both independently and as a mediating factor in the relationship between wages and endowments. Furthermore, the wage data used by Blum and Marshall are limited to manufacturing and to non-production and production workers, whereas from WIOD we have data on wages in all sectors by level of education (which is more readily comparable with endowments and with other evidence on skills and earnings).

Our paper also relates to a big and controversial literature on the effects of trade on wages in developed and developing countries (e.g. Wood, 1994, 1997; Anderson, 2005; Goldberg and Pavcnik, 2007; Harrison et al., 2011a; Leamer, 2012; Burstein and Vogel, 2012; Edwards and Lawrence, 2013). Over the past two decades it has become more widely accepted by economists that wages are substantially influenced by trade, but less widely accepted that this influence is of the sort suggested by HO theory, given scant empirical support for the HOS Stolper-Samuelson price-wage mechanism and evidence that greater openness has raised skilled wages in skill-scarce countries. Most studies of trade and wages, however, have been of changes over time within single countries, while we have panel data on wages of skilled and unskilled workers.

The rest of the paper is organised as follows. Section II sets out relevant theory. Section III introduces the WIOD data, with some descriptive statistics. Section IV presents the results. Section V concludes.

II Theoretical framework

For simplicity we work mainly with a 2×2 model of a single country, using the familiar 'hat' algebra of Jones (1965), extended briefly to many goods, but with only informal discussion of more factors and without introducing other countries (either bilateral trade or global general equilibrium). Support for the more general relevance of our analysis is provided by the work synthesised by Costinot and Rodrìguez-Clare (2014). The main purpose of this section is to set out the GHO model, but to clarify the properties of this model it is convenient to start by discussing a closed economy and the HOS model of an open economy.

Closed economy

Two factors, H (high-skilled workers) and L (low-skilled workers) produce two goods, B (biochemicals, which are H-intensive) and G (garments, which are L-intensive). Changes in the relative prices of the goods, p, which equal their factor costs, c, are related to changes in factor prices, w, by the zero-profit condition

$$(1) \quad \widehat{p}_B - \widehat{p}_G = \widehat{c}_B - \widehat{c}_G = (\theta_{HB} - \theta_{HG})(\widehat{w}_H - \widehat{w}_L)$$

where θ_{ij} is the share of factor i in the producer price or production cost of good j. A rise in the relative wage of skilled workers causes a rise in the relative cost and price of the skill-intensive good. Factor-market clearing requires

$$(2) \quad \widehat{v}_H - \widehat{v}_L = -\sigma_{BG}(\widehat{w}_H - \widehat{w}_L) + (\lambda_{HB} - \lambda_{LB})(\widehat{q}_B - \widehat{q}_G)$$

where the endowment of a factor is denoted by v, the output of a good by q, λ_{ij} is the share of the endowment of factor i used by good j, and

$$(3) \quad \sigma_{BG} = \sum_{j=B,G} [\lambda_{Hj}(1-\theta_{Hj}) + \lambda_{Lj}\theta_{Hj}]\,\sigma_j$$

is a weighted average of the elasticities of substitution in production between H and L for the goods, σ_B and σ_G. A rise (say) in the relative endowment of H must be matched by a rise in the relative demand for H, which can be achieved by a fall in the relative price of H that induces a rise in the H-intensity of the techniques used in producing both goods (the first rhs term in (2)) and/or by a shift in the composition of output towards the H-intensive good B (the second term).[2]

The final element of the closed-economy model is a demand function that links the relative quantities of goods sold to their relative prices

$$(4) \quad \widehat{q}_B - \widehat{q}_G = -\gamma_{BG}(\widehat{p}_B - \widehat{p}_G)$$

[2] Changes in the mix within sectors of goods of differing factor intensity are observationally equivalent to changes in technique.

where γ_{BG} is the elasticity of substitution in consumption between B and G. The effect of changes in endowments on factor prices in a closed economy can then be derived as

$$(5) \quad \widehat{w}_H - \widehat{w}_L = -\frac{1}{\sigma_{BG} + (\lambda_{HB} - \lambda_{LB})\,\gamma_{BG}\,(\theta_{HB} - \theta_{HG})}\,(\widehat{v}_H - \widehat{v}_L)$$

The first term in the denominator of the rhs ratio shows that endowments have more effect on factor prices if factors are less substitutable in production. The second term shows how changes in factor prices alter goods prices in ways that shift the composition of output in a direction that helps to absorb changes in endowments. The second term is the product of three elasticities: of relative goods prices with respect to relative factor prices $(\theta_{HB} - \theta_{HG})$, of relative outputs with respect to relative goods prices (γ_{BG}), and of relative factor use with respect to relative outputs $(\lambda_{HB} - \lambda_{LB})$. The lower the elasticities of substitution in production and consumption, and the smaller the difference in factor intensity between the goods, the more does a rise in the relative endowment of skilled workers in a closed economy depress their relative wage.

Heckscher-Ohlin-Samuelson

The key assumption of the HOS model is that, in an open economy, goods prices are no longer influenced by domestic demand, as in equation (4), but instead are determined by world prices and trade cost, requiring that

$$(6) \quad \widehat{c}_B - \widehat{c}_G = \left(\widehat{p}^*_B + \widehat{T}_B\right) - \left(\widehat{p}^*_G + \widehat{T}_G\right)$$

where p^*_j is the world price of good j and $T_j = \frac{c_j}{p^*_j}$ is the trade cost ratio (greater than unity if j is an import substitute and less than unity if j is an export good). With the usual assumption of iceberg (or other ad valorem) trade costs, the elasticities of demand for traded goods are infinite, which makes the ratio on the rhs of equation (6) zero (as if γ_{BG} had become infinite). Within a cone of diversification - a range of relative endowments bounded by the relative factor intensities of the two goods - variation in endowments does not affect relative factor prices, which are determined by

$$(7) \quad \widehat{w}_H - \widehat{w}_L = \frac{\widehat{c}_B - \widehat{c}_G}{\theta_{HB} - \theta_{HG}} = \frac{\left(\widehat{p}^*_B + \widehat{T}_B\right) - \left(\widehat{p}^*_G + \widehat{T}_G\right)}{\theta_{HB} - \theta_{HG}}$$

with the effect of changes in relative goods prices on relative factor prices magnified because

$(\theta_{HB} - \theta_{HG})$ is less than unity.

Equation (7) can illustrate the impact on factor prices of moving from autarky to trade, which involves both T_j's getting closer to unity. Consider for example a skill-abundant country, where in autarky $T_B < 1$ and $T_G > 1$ (the internal relative price of skill-intensive goods being low by comparison with the world): movements towards unity make $\widehat{T}_B > 0$ and $\widehat{T}_G < 0$, raising the relative wage of skilled workers. Similarly and more generally, across-the-board cuts in trade costs in an open economy raise the relative price of its abundant factor (given world prices, changes in which also alter factor prices).

The Jones algebra is not suited to the analysis of multiple cones of diversification, which is most conveniently (and familiarly) presented in a Lerner diagram (e.g. Schott, 2003; Xiang, 2007). In a model with two goods, countries with extreme endowment ratios lie outside the single cone and specialise in producing only one of the goods. In each region of specialisation, factor prices respond inversely to variation in endowments to a degree governed by the elasticity of substitution in production, σ, but are not affected by trade costs, though higher trade costs reduce absolute levels of factor prices by worsening the country's effective terms of trade (Markusen and Venables, 2007).

With many goods arranged in order of their relative factor intensity (still with only two immobile factors), there can be multiple cones, the countries in each of which produce a few goods of adjoining factor intensity (in the simplest case, only two). Between the cones can be regions in which only one good is produced. All other goods are imported from countries in other cones, implying a lot of trade and intense specialisation in each cone (rather than, as the label misleadingly suggests, diversification). Within each cone, factor prices are unrelated to endowments, but they are affected by trade costs. Across cones, factor prices vary inversely with endowments.

Given equation (7), the range of influences on relative factor prices in the HOS model is strikingly limited. Within a cone, variations in the internal demand for factors - due for example to differences in the size of nontraded sectors - affect only the composition of output and trade. Differences in technology affect factor prices, but only in certain ways: relative factor prices are affected by sector-biased technical differences, but not by factor-biased differences (though factor earnings will vary with factor quality).

Open-economy HOS models can easily accommodate traded intermediate inputs, though the immobile factor intensities of goods may then differ greatly from what they would have been in a closed economy. If the goods prices and trade costs in equation (6) refer to final goods, a fall in the cost of imported intermediates as a result of lower trade costs (or of lower outsourcing costs: Grossman and Rossi-Hansberg (2008)) has the same effect as technical progress in the sector concerned. For example, if the savings are greater in scarce-factor intensive sectors, the relative price of the scarce factor tends to rise, even if - as with outsourcing - the savings involve reduced use of the scarce factor.[3]

The HOS model can also accommodate more than two factors. Internationally mobile factors, which in the present paper are assumed to include capital, play a role similar to traded intermediates. Including more than two immobile factors - a third skill category, say, or land - leaves intact the HOS principle that factor prices are determined by world prices and trade costs, but complicates the details. The only general prediction is that, with equal numbers of goods and factors, for each factor there must be a good of which an increase in the price will lower the real return to the factor (Feenstra, 2003, p.70).

General Heckscher-Ohlin

In the GHO model of an open economy, the elasticity of demand for immobile factors is less than infinite. Equation (5) of the closed-economy model therefore becomes relevant again, but with the elasticity of demand in the goods market, instead of being γ_{BG}, becoming

(8) $\quad \epsilon_{BG}\delta_{BG}$

where ϵ_{BG}, the 'purchaser-price elasticity', measures the response of the relative sales and outputs of goods B and G to their relative purchaser prices, while δ_{BG}, the 'price-ratio elasticity', measures the response of the relative purchaser prices of goods B and G to their relative factor costs (and is less than unity).

(a) Purchaser-price elasticity

The relative sales of goods B and G by home producers depend on their relative prices, unlike HOS, where there is no such relationship. This relationship exists because in each sector there is a finite 'trade elasticity' that links the share of imports in domestic expenditure to the relative prices of imported and home-produced varieties.[4] Following Arkolakis et al.

[3]For the same reason that relative factor prices are not affected by factor-biased technical progress in the HOS model of a small open economy. The scarce factors released by the cost savings are absorbed by an increase in the relative output of the scarce-factor intensive sector.

[4]Strictly speaking, the trade elasticity refers to the ratio of imported to home-produced varieties rather than to the share of imports in expenditure, and is defined with respect to changes in trade costs rather than in prices more generally (Costinot and Rodrìguez-Clare, 2014, p. 201).

(2012) and Costinot and Rodrìguez-Clare (2014), we assume a CES utility function

$$(9) \quad C_j = \left[\left(C_j^H\right)^{\frac{\beta_j - 1}{\beta_j}} + \left(C_j^M\right)^{\frac{\beta_j - 1}{\beta_j}} \right]^{\frac{\beta_j - 1}{\beta_j}}$$

where C_j^H and C_j^M are composites of home-produced and imported varieties in sector j, and β_j equals one plus the trade elasticity.[5] This elasticity may reflect adjustments at either the intensive margin (more or less consumption of qualitatively different varieties, as in Armington, 1969, and Krugman, 1979) or the extensive margin (purchases of identical varieties from different countries or firms, as in Eaton and Kortum, 2002, and, partly, in Melitz, 2003). Relative expenditure on goods B and G from all sources depends on the relative prices of the B and G aggregates, in a way that is governed by a higher-level CES utility function

$$(10) \quad C = \left[\alpha C_B^{\frac{\gamma_{BG} - 1}{\gamma_{BG}}} + (1 - \alpha) \, C_G^{\frac{\gamma_{BG} - 1}{\gamma_{BG}}} \right]^{\frac{\gamma_{BG}}{\gamma_{BG} - 1}}$$

where α is a preference parameter and γ_{BG} as before is the elasticity of substitution between the goods, which is likely to be much lower than either β_B or β_G. The elasticity of the relative sales of domestic producers of B and G with respect to their relative purchaser prices is an average of γ_{BG} and the sectoral elasticities β_B and β_G. With (9) and (10) being CES, the average elasticity in any particular market can be written precisely, following Sato (1967), as a weighted harmonic mean, where the weights involve the shares of each of the goods in total expenditure and the shares of the country concerned in the sales of these goods in this market (Wood, 2012, section 2.1). A more tractable approximation to this average elasticity is the weighted arithmetic mean

$$(11) \quad \epsilon_{BG} = s_{BG} \gamma_{BG} + (1 - s_{BG}) \, \beta_{BG}$$

where s_{BG} is the country's average share of the sales of these goods in the market concerned and β_{BG} is an average of β_B and β_G. In the world market, s_{BG} is likely to be small, so ϵ_{BG} is close to β_{BG}. In the home market, however, domestic producers have a cost advantage, so that s_{BG} is likely to be big enough to make γ_{BG} matter, too. Home market shares vary among goods, depending on the country's comparative advantage, but for all goods depend

[5] The difference between the trade elasticity and β_j exists because the former refers to the value rather than to the volume of sales. The elasticities of substitution within the composites of home-produced and imported varieties are not necessarily the same size as β_j (Costinot and Rodrìguez-Clare, 2014, p. 244-6; Feenstra et al., 2014).

also on the country's international trade costs and policies.

The effect of relative purchaser prices on a country's relative sectoral outputs depends on its combined ϵ_{BG} across all markets, which is an average of ϵ_{BG} in its home and export markets, weighted by its shares of total sales in each market. This combined elasticity decreases with the height of a country's international trade costs, averaged across all goods - or equivalently increases with its openness to trade - for two reasons. Higher trade costs raise s_{BG} and thus lower ϵ_{BG} in its home market.[6] They also reduce the share of exports in its output and so the weight in the combined ϵ_{BG} of the higher ϵ_{BG} in the world market (where s_{BG} is small).

(b) Price-ratio elasticity[7]

The price-ratio elasticity is less than unity because the purchaser price of each good is the sum of its internationally immobile factor costs (IFC) and an other-cost wedge (OCW) that includes trade costs, purchases of traded intermediates and payments to mobile factors (not least, by assumption in this paper, capital). OCWs are often big, relative to IFCs, and usually do not vary in proportion to IFCs. Denoting the OCW per unit of output of good j by t_j, the IFC of good j by c_j (as before), and defining $\tau_j \equiv \frac{t_j}{c_j}$, the price-ratio elasticity δ_{BG} is determined approximately by

$$(12) \quad \delta_{BG} = \frac{1 + \eta_{BG}\tau_{BG}}{1 + \tau_{BG}}$$

where τ_{BG} is the geometric mean of τ_B and τ_G, and η_{BG} is the elasticity of $\frac{t_B}{t_G}$ with respect to $\frac{c_B}{c_G}$.

To understand equation (12), consider the expression $\frac{1}{1+\tau_{BG}}$, which is what (12) would become if $\eta_{BG} = 0$ and is the average share of IFCs in the purchaser price. The smaller this share, as a result of a larger τ_{BG}, the smaller the effect on relative purchaser prices of a proportional change in relative IFCs (just as, for example, with c_j half of p_j, a 10% rise in c_j, with no change in t_j, would raise p_j by only 5%). However, insofar as relative OCWs vary in proportion to relative IFCs, for example if some trade costs are ad-valorem, η_{BG} will be positive, tending to increase δ_{BG} (and if η_{BG} were unity, as if for example OCWs consisted only of ad-valorem trade costs, δ_{BG} would be unity, too).

[6] More precisely, s_{BG} depends on the average across the two sectors of the proportional cost disadvantage of foreign suppliers relative to home suppliers and on the average 'trade elasticity' ($\beta_j - 1$).

[7] For a fuller exposition of the next two paragraphs, see sections 2.2 and 2.3 of Wood (2012).

(c) Relative wages

Extending the model to include n goods, indexed by j and with good 1 as the numeraire (but still with only two factors), relative wages are determined by

(13)

$$\widehat{w}_H - \widehat{w}_L = -\frac{1}{\sum_{i=1}^{n}\left[\lambda_{Hj}\left(1-\theta_{Hj}\right)+\lambda_{Lj}\theta_{Hj}\right]\sigma_j + \sum_{j=2}^{n}\left(\lambda_{Hj}-\lambda_{Lj}\right)\epsilon_{j1}\delta_{j1}\left(\theta_{Hj}-\theta_{H1}\right)}\left(\widehat{v}_H - \widehat{v}_L\right)$$

Relative factor prices in the GHO model are affected by variation in endowments, even in open economies producing more than one good, because the ϵ_{j1}'s are finite. Changes in output mix require changes in relative goods prices (because demand is less than infinitely elastic), which in turn require changes in factor prices, whose size is amplified because the δ_{j1}'s are less than unity, muting the effect on relative purchaser prices of changes in relative IFCs. The inverse relationship between endowments and factor prices is continuous (rather than stepped as in multi-cone HOS), but tends to be less steep in a more open economy. As explained above, lower international trade costs increase the average ϵ_{j1}. Lower per-unit international trade costs also raise the δ_{j1}, increasing the average price-ratio elasticity.

In the GHO model, the effects of changes in world prices and trade barriers are in the same directions as in HOS, but the channels of influence are different. With different varieties of goods, the direct links in equation (7) cannot apply. Changes in the domestic prices of foreign varieties as a result of changes in foreign costs or trade barriers alter factor prices indirectly, by shifting the demand for substitute domestic varieties.

In GHO, traded intermediates are part of the OCW, since they are a component of costs that is common to all countries and does not vary with factor endowments, so they tend to reduce price-ratio elasticities. As in HOS, however, trade in intermediates also affects the immobile factor intensities of goods, since countries tend to import intermediates in which they have a comparative disadvantage and thus make more intensive use of their abundant factors. Trade in intermediates will show up in equation (13) as absolutely larger $\theta_{Hj} - \theta_{H1}$ and $\lambda_{Hj} - \lambda_{Lj}$, pulling in the opposite direction to the fall in δ_{j1} caused by the higher ratio of OCWs to IFCs. The net effect in GHO of more trade in intermediates, as of other sorts of trade, is likely to be less sensitivity of relative factor prices to variation in relative endowments.[8]

In GHO, payments to mobile factors are also part of OCWs, reducing price-ratio elasticities. Changes in the world prices of mobile factors can alter the prices of immobile factors through sector-biased channels and factor-biased channels, the latter depending on substitutability and complementarity between the factors. With more than two immobile factors, substitutability and complementarity also influence the effects of variation in immobile factor endowments on factor prices, since the effect of a change in the supply of one factor on the relative prices of other factors depends in GHO on how it affects the relative demand for them.

In summary, a key property of the GHO model (as implied also by Costinot and Rodrìguez-Clare, 2014, p. 222-3, and by Burstein and Vogel, 2011) is that factor prices in an open economy are determined not directly by world prices and trade costs (as in HOS) but by the balance of supply and demand in the country's factor market, which is indirectly influenced by trade but depends also on endowments. Factor prices can thus also be affected by other determinants of relative factor demands, such as autonomous changes in the domestic consumption mix or factor-biased technical change. GHO can therefore explain not only why factor prices vary with endowments in open economies, without the awkwardness of HOS cones, but also why the sensitivity of factor prices to endowments varies among countries and time periods, being greater where barriers to trade are higher and other-cost-component wedges are larger.

III Data and descriptive statistics

The data used in this paper are drawn from the World Input-Output Database (WIOD), a recent resource described by its creators in Timmer (2012). It offers a single and consistent source of global data on output, trade, factor use and factor prices, putting its users at an advantage over the authors of earlier studies reviewed in section I, who have had to put together their data from several different sources. WIOD too is compiled from different sources, but its compilation was unusually thorough.

The core of WIOD is annual input-output tables for 1995-2009 that connect 40 countries - 27 members of the European Union and 13 other major economies, in total accounting for 85% of world GDP, plus a composite rest of the world. The main diagonal of each year's

[8] However, where the saving in costs from using imported rather than home-produced intermediates is small, it is possible, even with only two countries, that the net effect could be greater sensitivity. With more than two countries, it is also possible that trade in intermediates could reduce rather than increase $\theta_{Hj} - \theta_{H1}$ and $\lambda_{Hj} - \lambda_{Lj}$, which would guarantee greater sensitivity. Details available on request.

input-output table consists of tables of intermediate flows within each country among 35 industries, with off-diagonal tables showing trade in intermediates among all the countries and other columns covering final demand in each country, supplied both domestically and from imports. All the flows in the table are values at basic prices, but information is also available on trade and transport margins for internal and international transactions. Data on the volumes and prices associated with value flows are not available.

Among WIOD's auxiliary tables are socio-economic accounts providing (among other things) information on the levels of employment and wage bills of three skill categories of worker in every country, industry and year. Skill is measured by length of schooling, following the International Standard Classification of Education (ISCED). 'Low skilled' workers are ISCED categories 0, 1 and 2 (everything below completed upper secondary). The 'medium skilled' are ISCED categories 3 and 4 (complete upper secondary and some tertiary, but below a college bachelor's degree), and the 'high skilled' are in categories 5 and 6 (a college bachelor's degree and above).

WIOD wage and employment data were assembled from national labour force surveys and censuses, which have not previously been collated in this form. We used these data to derive wage rates - the form in which wages were typically reported in the sources. Though these data are by far the best available, their accuracy is open to doubt, especially in poorer countries with large numbers of self-employed workers, for whom wages comparable to those of employees had to be imputed. In some countries, gaps had to be filled by using data from other similar countries. In addition, of course, the quality of schooling varies widely.

Another reservation about using WIOD to analyse relationships between trade and factor prices is that the countries in the dataset are atypically large and therefore trade atypically little. Though these countries account for 85% of world GDP, they are only about one-fifth of all the countries in the world and thus on average about four times as big as a world average country. With more small countries, our results might have been different.

A first glance at the data is in Figure 1, which plots the relative wage of skilled workers across countries against relative endowments of skilled labour in four years spanning the full period. As in Timmer et al. (2014), we aggregate the three WIOD skill categories into two by combining 'low-skilled' and 'medium-skilled' into 'unskilled', who are thus all those with less than a college degree.

[Insert Figure 1 here]

Though purely descriptive, the plots show a negative linear relationship: in countries with relatively more skilled workers, the relative wage of skilled workers tends to be lower. This relationship is however far from perfect, with wide dispersion among countries with low skill endowments: India, Brazil and Indonesia have the highest skill premia, but China one of the lowest, especially in the early years. Moreover, the negative relationship with wages is not consistent at intermediate endowment levels.

Other variables are used in the empirical analysis to control for the level of economic development, additional factors of production and their composition, and labour market institutions. GDP per capita (in 2005 constant US$) and population are sourced from the World Development Indicators, while data on human capital are from the Barro and Lee (2013) database. To control for the stringency of labour market institutions, we use data on unionization and collective bargaining from the ICTWSS database (Visser, 2013), indicators on Employment Protection Legislation from the OECD (2013) and the "Labour Freedom" index from the Heritage Foundation.

Table 1 reports summary statistics for the variables used in later regressions. Table A1 in the Online Appendix is a correlation matrix for the main variables.

[Insert Table 1 here]

IV Empirical analysis and results

The confrontation of theory with data in this section proceeds as follows. In Part A, we scrutinise the relationship between relative wages and relative endowments. We investigate whether the negative slope in Figure 1 is sensitive to the inclusion of control variables or explained by variation in workers' quality. We also check in the cross-country data for the existence of multiple cones of diversification. We then test for factor price insensitivity to endowment variations within countries over time.

In Part B, we investigate the effects on relative wages of the height of barriers to trade, with special reference to how these effects vary with endowments, since HO theory predicts that the usual effect of greater openness will be to benefit a country's relatively abundant factor. In Part C, we examine whether variation in price-ratio elasticities, as well as variation in openness, helps to explain variation in the strength of the consistently inverse relationship between relative wages and skill endowments. In Part D, we estimate changes in skill abundance as predicted by different versions of the GHO model and assess how they match with changes as observed in the data.

A Wage-endowment relationship

All WIOD countries in all years engaged substantially in trade and thus were open in the sense relevant to the HOS model, which is compatible with high trade barriers and nontraded sectors.[9] So the HOS prediction of insensitivity of factor prices to variation in endowments within cones should apply to them. Our first simple test of this prediction is to regress the relative wage of skilled workers on relative endowments of skilled labour, with both variables in logs to reduce the influence of the outliers in Figure 1.

Figure 2 shows the results of estimating a separate elasticity for each year across all 40 countries, reported in the upper panel with 95% confidence intervals. The results show that a 10% increase in the relative supply of skilled workers is associated with roughly a 3% decrease in their relative wage. The elasticity is negative and significant in all years, and fairly constant over time, though more precisely estimated in later years, as shown in the lower panel by the rise in the adjusted R^2, confirming the impression given by the regressions in Figure 1.[10] Unlogged regressions in Figure A1 in the Online Appendix generate a similar pattern of significantly negative coefficients, but fit less well.

[Insert Figure 2 here]

We perform a number of robustness checks on this negative cross-country relationship between relative wages and endowments, using the natural logarithm of the average of each variable over time for each country. Regression results are reported in Table 2. The estimated elasticity of average wages with respect to average endowments is close to the average of the yearly elasticities in Figure 2 (see column (1)).

[Insert Table 2 here]

As a first set of checks, we modify our definition of 'unskilled', a category that includes a wide range of schooling levels - from none to some tertiary education. In column (2), we control for the share of 'low' in the unskilled (low plus medium) aggregate and find that the skilled wage premium is higher where the unskilled on average have less education - though the wage-endowment elasticity is much the same as in column (1). The low-skill category itself is broad - from no education to incomplete upper secondary - and the mixture varies a lot across countries. In column (3), adding the share of low-skilled workers without any

[9] In 2000 five of them - China, Estonia, India, Malta and Russia - were still classified as "closed" on Sachs-Warner criteria (Wacziarg and Welch, 2008), but even they were not autarkic.

[10] Dropping China increases the fit almost equally in all years.

education (computed from the Barro and Lee, 2013 database[11]) again shows that the skilled wage premium is higher where unskilled workers have less education - and the coefficient on the share of low-skilled hours drops to zero, suggesting that the proportion of people with no education is the main driver of this composition effect - though again the coefficient on the endowment variable is virtually unchanged. In column (4), we shift medium-skilled workers from the 'unskilled' to the 'skilled' category, which weakens the coefficient and worsens the fit, supporting our assumption that medium-skilled workers (those with complete upper secondary education, who on average in WIOD countries are about half of labour supply) are more substitutable for low-skilled workers than for high-skilled ones (with a college degree or more).

As mentioned earlier, endowments of other immobile factors of production, if they are substitutes or complements for skill or labour, can affect the relationship between the relative wages and relative endowments of skilled and unskilled workers. One such factor is land, which is likely to be complementary to unskilled labour, so in column (5) we add the ratio of agricultural land area to unskilled labour, yielding a coefficient with the expected (negative) sign, but insignificant and not altering the elasticity of w with respect to v in the first row. Following Wood (1994), we treat capital as a mobile factor and in column (6) add an estimate of the cost or rental rate of capital, derived imperfectly from WIOD as the ratio of non-wage value added to the value of the fixed capital stock. The usual assumption of capital-skill complementarity would predict a negative coefficient (with more expensive capital reducing the demand for skill), but it is positive, though again insignificant and with no change in the w-v elasticity. The same is true when both land and the cost of capital are included in column (7).

The inverse relationship between skilled wages and skill abundance could be driven by other country-level factors. Given the limited sample size, we focus on economic development and labour market institutions as two plausible candidates. Specifically, in column (8) we include GDP per capita and the Labour Freedom index (the *labmkt* variable, on which data are available only from 2005). The negative and significant relationship between the skilled wage premium and skill abundance remains, but becomes smaller, because the skill endowment variable is correlated with per capita income. More rigid labour markets (a higher minimum wage and more regulations on hiring and firing) tend to reduce the wages of skilled workers relative to unskilled workers. Column (6) confirms this last result with

[11] We take the ratio (in logs) of people with no education to people with "attained" (not completed) secondary education at most - i.e. the most comparable category to the 'unskilled' definition from WIOD data. While WIOD data on labor endowments are in hours worked, the Barro and Lee data are in number of people regardless of their employment status.

another measure of labour market rigidity - the share of workers unionised. Similar results were obtained with other measures of labour market institutions.[12]

The negative and significant cross-country relationship between relative skilled wages and endowments thus seems robust. As a final check, we regress logged relative wages in each sector on logged country-level endowments, including sector dummies to pick up wage differences across sectors that are common to countries. This specification reduces the risk of bias caused by national average relative wages being affected by variations in sectoral employment shares due to differences in endowments. The cross-section results reported in Figure 3 are similar to those in Figure 2, suggesting that this potential source of bias is of minor importance, and these results survive the same set of robustness checks as in Table 2 (reported in Table A2 of the Online Appendix).

[Insert Figure 3 here]

Cross-country differences in factor quality or productivity

These negative cross-country coefficients are at face value inconsistent with the textbook one-cone HOS model, but could be reconciled with HOS in two possible ways. One possible reconciliation, to be explored shortly, is that countries are in different cones of diversification. Another, in light of evidence that the effectiveness of schooling varies enormously, is that the observed cross-country differences in the relative wages of workers with different amounts of schooling are an illusion arising from differences in the quality or productivity of skilled or unskilled workers. For example, the quality of basic education may vary more than that of higher education because expansion of higher education provides a larger supply of qualified teachers at lower levels.

Adjusted for quality, the relative wages of skilled and unskilled workers might thus vary much less across countries than the unadjusted wage data suggest. To test this hypothesis, we apply to our cross-country data the ingenious method that Bernard et al. (2013) developed to test for factor price equalisation across US states. They allow for locality- and sector-specific differences in factor quality by using data on relative wage bills (rather than wage rates), in which, assuming cost minimisation, unobserved factor qualities cancel out. Our application of this method is reported in section AIII of the Online Appendix, partly to save space but

[12] We use other proxies for labour market institutions as in Freeman (2007). The share of employees covered by collective wage bargaining has the same effect as unionisation. We also used the indices of restrictiveness of Employment Protection Legislation from the OECD, though these are available for only 33 countries: the wage-endowment elasticity slightly decreases and loses significance; stricter regulation of regular contracts lowers the relative wages of the skilled, while stricter regulation of temporary contracts puzzlingly has the opposite effect.

mainly because our results, like those of Bernard et al. (2013) for the US, strongly reject the hypothesis of relative factor price equalisation.

Adjusting for differences in quality in this way, the inverse relationship between relative wages and relative endowments becomes even stronger, both across countries and over time. This implies that the quality of more educated workers relative to the quality of less educated workers rises (rather than, as hypothesised, falls) as the proportion of more educated workers in the labour force increases. However, because the size of the quality adjustment depends on assumptions about the size of elasticities of substitution across factors, for the rest of this paper we revert to using unadjusted wages.

Multiple cones of diversification

The negative elasticity of relative wages with respect to relative endowments across countries reported above could be reconciled with a HOS model with multiple cones of diversification. In such a model, countries in different endowment ranges specialise in subsets of sectors of different skill intensity, and have different relative wages, though within each cone relative factor prices do not respond to variation in endowments.

Kiyota (2011, 2012a) finds evidence of multiple cones of diversification in Japan, looking at both output mix and factor prices. In this paper, we focus on factor prices, leaving for future iterations of our work a check for matching patterns of output specialisation. Further, we follow Schott (2003) in assuming that there are no regions of complete specialisation, where the response of factor prices to endowments is downward-sloping, in between the factor price 'plateaux' of the cones. We adopt this simplifying assumption mainly because of the fewness of our countries (only 40), which makes it statistically difficult to identify many different cones and regions.

Given these assumptions, we take the multiple-cone HOS prediction at face value and estimate in the within-country averaged data:

$$(14) \quad \ln(w_c) = \sum_{d=1}^{D} \beta_d I_d\{\ln(v_c) > \overline{v}_d\} + \varepsilon_c$$

where $\overline{v}_d$ is the threshold value of $\ln(v_c)$ that identifies the dth interior knot. The term $I\{\cdot\}$ denotes an indicator function equal to one if the expression in brackets is satisfied and to zero otherwise. In a model with N cones of diversification, $N-1$ interior knots are estimated. We search for the location of the interior knots by gridding over values of $\ln(v_c)$ from its minimum to its maximum and using a grid interval of 0.2.[13] We test for a maximum of four cones (with three interior knots) and choose the set with the lowest Akaike Information Criterion (AIC).

The results are consistent with the existence of multiple cones of diversification in the sense that specifications with more cones fit the data better - the best fit is for the specification with four cones. Figure 4 plots a scatter of relative skilled wages against relative endowments of skilled labour (averaged over time and in logs), with a horizontal line showing the predicted wages from the best-fitting four-cone model and vertical lines showing the knots, which correspond to college-educated labour supply of 6%, 12% and 20%. China is alone in the lowest cone, excluding which relative wages decline across cones with relative endowments, as predicted by theory. We obtained similar results using sector-level wage data (results available upon request).

[Insert Figure 4 here]

The multi-cone HOS model allows us to analyse the "paths" of development that countries follow while moving across cones of diversification (Leamer, 1987). While a fuller treatment is beyond the scope of this paper and left for future work (see Deardorff, 2001; Kiyota, 2012b for theoretical studies), here we empirically explore the evolution of skill abundance within countries holding the values of the cones boundaries (the knots) fixed at those estimated in Figure 4. As shown in Figure A2 in the Online Appendix, skill abundance rises between 1995 and 2009 in all countries except Mexico, where it stays constant. Half of the countries move to a more skill-abundant cone during the sample period, with most of the action occurring in the last years. This type of analysis however does not allow the boundaries of the cones to vary over time. We thus replicate the estimation of equation (14) year by year. Figure A3 in the Online Appendix shows the results in a way similar to Figure 4, but now highlighting countries that moved up or down the spectrum of skill-abundance cones relative to the previous period in the figure. The four-cone specification is the preferred one in all years and the position of the knots does not vary much over time. Most of the movements up to a more skill-abundant cone occur between 2005 and 2009, where India and Indonesia remain in the second least skill-abundant cone while other countries join more 'skilled' cones.

These results give support to the multi-cone version of HOS not only because the fit with four cones is much better than with one cone (where log relative wages are regressed on a constant term) but also because the four-cone specification fits better than the linear

[13] We experimented with even smaller intervals of 0.1 and 0.05. The optimal number of cones stays the same, although countries' location across cones varies slightly. Reducing the interval can only support the finding of additional cones.

model above which regressed log relative wages on log relative endowments. However, the fact that the statistical fit improves steadily with the number of cones suggests that the linear specification may approximate the true relationship between relative wages and endowments - in other words, that the true relationship is continuously declining rather than stepped. We could not test this hypothesis by increasing the possible number of cones since, as is already evident with four cones, the role of single-country observations would become even greater.

Factor price insensitivity in time series

The relevance of the multi-cone HOS model can be further assessed by analysing the time-series dimension of the data, since if the model is accurate relative wages should differ among countries, depending on which cone they are in but should not be sensitive to changes over time in endowments within each country (so long as these changes are not large enough to move a country between cones).

We thus pool the data across years (1995 to 2009) and exploit within-country variation over time in wages and endowments. Specifically, we estimate:

$$(15) \quad \ln(w_{c,t}) = \beta \ln(v_{c,t}) + \alpha_c + \phi_t + \varepsilon_{c,t}$$

The country fixed effects α_c control for all time-invariant country-specific characteristics so that the coefficient of interest β relates variation over time in wages in each country to variation over time in its endowments.

Table 3 reports the within-country estimates of equation (15) and robustness checks. On an annual basis, relative wages move inversely with relative endowments, although the effect is imprecisely estimated and much smaller than across countries (column (1)), perhaps because there is little yearly variation in endowments. This result survives our experiments with skill categories (columns (2)-(4))[14] and allowance in columns (5) and (6) for variation over time in land abundance (slight) and in the cost of capital. However, the coefficient on the cost of capital is significantly and substantially positive, which is consistent with the finding of Timmer et al. (2014), using WIOD data, that the shares of both skilled wages and capital in value added rose in most countries during this period. The controls in column (7) for economic development and unionisation slightly reduce the wage-endowment elasticity,

[14] The Barro and Lee data are available in 5-year intervals, starting from 1995 in our sample. We use 2010 values in 2009. The alternative measure of the skill ratio in column (4) performs more strongly, relative to our standard measure, in these time-series estimates than in the cross-section estimates in Table 2.

and suggest a positive association within countries over time between the rate of economic growth and changes in the skilled wage premium.

[Insert Table 3 here]

Two possible concerns about the panel specification in (15) and its extensions are that endowments (and hence wages) do not vary much on a yearly basis and that the effects on wages of changes in endowments may be lagged. To allow for these concerns, we also follow the lead of Blum (2010) and estimate a panel specification with time spans of different lengths. Variables are transformed into annualized rates of change: $\widehat{x}^{l}_{c,t} \equiv {}^{\left(\ln(x_{c,t}) - \ln(x_{c,t-l})\right)}\!/_{l}$.[15] The regression becomes:

$$(16) \quad \widehat{w}^{l}_{c,t} = \beta \widehat{v}^{l}_{c,t} + \alpha^{l}_{c} + \phi^{l}_{t} + \varepsilon^{l}_{c,t}$$

where country dummies α^{l}_{c}'s control for country-specific trends in relative wages and time dummies ϕ^{l}_{t}'s for global trends. Like Blum (2010), we experiment with different lengths, l, of the time window. Longer windows might provide greater statistical power in identifying our coefficient of interest, since national factor endowments change only slowly. The HOS model, however, suggests a different pattern (Leamer and Levinsohn, 1995), with the initial impact on factor prices of a change in endowments fading away as the output mix adjusts, which would predict larger coefficients with shorter windows and a zero coefficient only with a long enough window.

Table 4 reports the estimates of the benchmark specifications together with a selection of robustness tests. As we extend the time window, the negative relationship between wages and endowment relationship becomes stronger and more precise. For five-year changes, a 10% increase in the relative endowments of skilled labour is associated with a 2% decline in the skill premium (column (1)). This estimate is unchanged when we control for changes in the share of low-skilled among the unskilled (column (2)), but is much weaker when medium-skilled workers are put into the 'skilled' category (column (3)), as in the cross-country estimates. The relationship remains negative and significant after controlling for changes in other factors (column (4)) and GDP per capita and union membership (column (5)).[16] The wage-endowment relationship becomes even stronger with 10-year changes: a 10% increase in relative skill abundance is correlated with a 3% decline in the skill premium

[15] This transformation is approximate, but exact annualisation produces statistically identical estimates.

[16] Results from the other robustness tests are available upon request. Note that the no-schooling (v_{ns}) variable has variation in only two periods with five-year changes and one period with 10- and 14-year changes. The *union* membership variable was selected among the labor market indicators simply because of greater data availability.

(column (6)), even after the robustness checks in columns (7), (8) and (10).[17] For windows beyond 10 years, however, the coefficient becomes absolutely smaller, being insignificant after 11 years and virtually zero for the longest possible window of 14 years (columns (11) to (15), the significant coefficient in the last of which arises from restriction of the sample). This pattern is qualitatively consistent with the Leamer and Levinsohn (1995) HOS prediction, but the implied output adjustment lag is implausibly long. Also, over a decade or more other forces may be at work - supply-side responses and induced technical change (Leamer, 2012, p. 109; Blum, 2010).

[Insert Table 4 here]

Table 5 reports corresponding results, in levels and changes, using sectoral wage data, as in Blum (2010), with suitable modifications of the panel specifications in equations (12) and (13), including the addition of sector fixed effects.[18] There is now no apparent effect on wages of changes in endowments in the levels specification or with five-year changes, a significant negative effect in the ten-year window (but smaller than in Table 4), and again no significant effect in the longest possible window.

[Insert Table 5 here]

In some respects, these results are similar to those of Blum (2010), who finds an inverse relationship between changes over time in relative supplies of skilled workers and their relative wages and also that this relationship is stronger over a decade than in shorter windows. Unlike us, though, Blum finds that the negative relationship becomes even stronger beyond a decade - which he attributes to induced scarce-factor-biased technical progress. One possible reason for the difference in results is that Blum's data cover only manufacturing, while ours cover all sectors: however, this explanation can be rejected by repeating our calculations for manufacturing only, as in Table A3 in the Online Appendix, with results that are similar to those for the whole sample using country-level (Tables 3 and 4) and sector-level (Table 5) wages. Blum's data however differ from ours also in their time period, country coverage and measure of skill (his is occupational - non-production and production workers - while ours is educational).

Other studies of changes over time in individual countries have generated similar results - increases in the relative supply of more educated workers apparently lowering their relative

[17] With 10-year changes, the *EU* dummy becomes collinear with the *union* variable since the only two countries that changed EU status in the ten-year windows, Bulgaria and Romania, have variation in *union* membership only in the same year of EU accession.

[18] Results of the different robustness test using sector-level wage data are available upon request.

wages (e.g. Robbins, 1996; Katz and Murphy, 1992). There is thus little empirical support for the existence of factor price insensitivity, even within individual countries over time. These findings could in principle be reconciled with HOS by supposing that they arise from movement of countries across cones, but this would imply the existence of cones so numerous and narrow as to make the HOS analysis of outcomes within each cone of very little relevance as a description of reality.

B Influence of trade barriers

A core prediction of any HO model - HOS or GHO - is that an across-the-board reduction of a country's barriers to trade tends to raise the price of its abundant factor relative to that of its scarce factor. Beyond simple but unrealistic comparisons between autarky and free trade, this prediction needs careful definition: 'across-the-board' should refer to uniform reductions of barriers that were initially uniform across sectors, and in effective rather than nominal terms, a pattern of change that is also unlikely to occur in reality. So there will be exceptions, even if HO theory is valid, but the presumption is that on average this prediction will apply with any broad measure of openness to trade.

To measure openness, we follow many other scholars in using ratios of total trade (exports plus imports) to output. An advantage of WIOD is that the denominator can be gross output, matching the gross measure of trade in the numerator, rather than, as in most earlier studies, GDP. However, a drawback of country-level trade/output ratios as a measure of openness is that they tend to be substantially lower in large countries than in small countries, not mainly because of higher trade barriers (though longer internal distances do add to trade costs), but because of more potential for realising economies of scale internally and wider diversity of natural resources.[19] To alleviate this size bias, we measure openness by the residuals of a cross-section regression of the trade/output ratio on population (both in logs). The results to be reported below would be similar if we did not log the measure of openness.[20]

To test the core prediction of HO theory, we interact our openness measure with relative endowments, expecting to find a positive coefficient on the interaction term. In a repeated

[19] Gravity-based measures of trade barriers (e.g. the Constructed Home Bias by Anderson and Yotov, 2010), while properly controlling for internal distance, are also strongly (and inversely) correlated with country size.

[20] By using the exponential of the residuals from the regression of the trade/output ratio (in logs) on population (in logs). We also get similar results if we do not adjust the openness ratio for country size but instead include (as additional variables in the regression) country size and its interaction with endowments. Results are available upon request.

cross-section framework, our benchmark specification is:

$$(17) \quad \ln(w_c) = \alpha + \beta_1 \ln(v_c) + \beta_2 \ln(o_c) + \beta_3 \left(\ln(v_c) \times \ln(o_c)\right) + \varepsilon_c$$

Where $\ln(o_c)$ is our openness measure (in logs) - the coefficient on which should in HO theory be zero for a country with world average endowments. Standard errors are computed using the bootstrap to control for the generated regressor. Results are shown in Table 6 for within-country averaged data. In column (1), relative wages are regressed on the openness variable alone. There appears to be no unconditional relationship between trade openness and relative wages, which is consistent with HO theory, but inconsistent with the often-suggested existence of other mechanisms by which greater openness could raise the relative wages of skilled workers (Wood, 2002; Harrison et al., 2011b; Burstein and Vogel, 2011).[21]

Controlling for factor abundance in column (2) makes the wage-openness elasticity larger and significant at the 10% level. The wage-endowment elasticity is similar to that estimated without controlling for openness in Table 2. Column (3) reports on our baseline specification in equation (14). The positive sign of the interaction coefficient matches the HO prediction, though it is not statistically different from zero, perhaps because of the crudity of our measure of the height of trade barriers. Further, this result masks variation in the effect over time. In Figure A4 in the Online Appendix, we show the estimated openness-endowment interaction coefficient and its 95% confidence interval in all years. While the coefficient is always positive, only in seven years is it also statistically significant at the 10% level or better.

[Insert Table 6 here]

The other columns in Table 6 take the specification in column (3) through much the same robustness tests as used earlier. In columns (4) and (5) we include the low-skill and no-schooling shares, both alone and interacted with the openness variable, which leaves the openness-endowment interaction positive and similar in size to column (3). The same is true when we control for land and the cost of capital (columns (6) and (7)). In column (8) we include GDP per capita and the Labor Freedom index, which makes the openness-endowment coefficient larger and significant, suggesting a confounding effect of economic development. When we interact the openness variable with these other regressors in column (9), the endowment-openness coefficient is less precisely estimated but still larger than in the baseline

[21] However, the decline over time in the size of the openness-endowment interaction coefficient in Figure A4 might reflect a greater rise in the relative importance of non-HO forces in skill-scarce (developing) countries than in skill-abundant (developed) countries.

specification. Similar results are obtained with union membership as the measure of labour market institutions (columns (10) and (11)).[22]

A possible explanation of the decline in the size of the interaction coefficient over time (Figure A2) is the increasing inaccuracy of this crude measure of the height of barriers in an ever-more-integrated world. One much-remarked feature of changes in trade over this period is the increasing share of trade in intermediate goods. A full investigation is beyond the scope of this paper, but as a preliminary step, in Table A4 in the Online Appendix, we split the openness measure between trade in final goods and trade in intermediates and use the resulting variables in our baseline specification.[23] Openness in final goods trade has a negative but not significant effect on the relative skilled wage, while the effect of openness in intermediate goods trade is practically zero (column (1)). Similarly, when we interact both measures with skill endowments, we find a positive and significant interaction effect only for openness in final goods trade (column (3)). These estimates should be interpreted with caution given the high correlation (rho = 0.7) between the two measures of openness.

As another measure of openness, we also tried the Overall Trade Restrictiveness Index (OTRI) estimated by Kee et al. (2009), following the theory of Anderson and Neary (2005), which is the uniform tariff rate that if imposed instead of the existing structure of protection would not alter a country's total imports. We follow the methodology of Kee et al. (2009) and use their estimates of the import demand elasticity and ad-valorem equivalent (AVE) of non-tariff measures, both for years around 2008-2009. Applied MFN tariffs and product-level imports are sourced from TRAINS and COMTRADE. This measure should capture only policy-related and border barriers to trade. Furthermore, limited tariff data and the use of year-specific elasticities and AVE estimates are likely to introduce measurement error. Bearing these caveats in mind, columns (4) to (6) of Table A4 report wage-openness regressions using the OTRI indicator (in logs). Interestingly, the indicator is strongly correlated with skill abundance (rho=0.5), which explains why its coefficient is positive and significant in column (4) but loses significance when we control for skill abundance in column (5). The sign of the interaction coefficient in column (6) is consistent with theory - protection is bad for skilled workers in skill abundant countries - but it is poorly estimated.

As a further robustness check, we regress sector-level relative wages on country-level endowments, openness and their interaction, including sector fixed effects. Figure A5 in the Online Appendix graphs the estimated interaction effect by year. The results are weaker

[22]Results (available upon request) are qualitatively similar when using the collective bargaining variable or the EPL indicators.

[23]Precisely, the two measures are the (within-country average of) residuals of two separate regressions of the final goods trade ratio to output and intermediates trade ratio to output onto population (all in logs).

than at the country level, with the coefficient on the interaction effect being positive and significant at the 10% level only in 2004. We also relate sector-level wages to sector-level openness, expecting to find, as predicted by HO theory, no relationship, because skilled and unskilled labour are mobile across sectors. On average over the period the elasticity is positive and just significant, but very small - results are available upon request.

The analysis so far has been of cross-country relationships, but in principle the core HO prediction about the effect of openness on relative wages applies also within countries over time. To test the prediction in this dimension we extend the panel regression specifications in equation (12) to include an interaction term between openness and relative endowments as well as openness alone, with results shown in Table 7. This specification controls for time-invariant country-specific forces that might influence both relative wages and openness conditional on factor abundance.

Columns (1) and (2) show that relative wages do not vary with trade openness over time. The coefficient on the endowment-openness interaction in column (3) is positive and statistically significant, as predicted by HO theory. Moreover, this interaction effect is of similar size to the one in the cross-section regression (see Table 6), but more precisely estimated. To better assess the importance of the openness effect, in Figure 5 we plot the elasticity of relative wages with respect to openness against relative factor endowments (in logs), using the estimates in column (3). The results imply that if a relatively skill-poor country like Turkey, where skilled labor was only 6% of unskilled labor in 1995, were to raise its skill endowment ratio to around 22%, its wage-openness elasticity would go from a substantial -0.7 to zero. Importantly, the openness elasticity turns positive and significant when countries become relatively skill-abundant.

The positive and significant coefficient on the interaction term is confirmed in Table 7 by several robustness tests that control for additional determinants of relative wages both independently and interacted with the openness variable. Specifically, results are mostly unchanged when controlling for skill composition (columns (4) and (5)), land and capital (columns (6) and (7)), again exposing the positive relationship over time between the skill premium and the rent of capital, and GDP per capita and union membership (columns (8) and (9)).[24] Table A5 in the Online Appendix shows that with longer time windows in the changes specification, both the interaction coefficients and the openness coefficients are smaller and less significant, except with the longest possible (14-year) time window.

[24]Results using the OTRI (not shown) are not robust probably because of substantial measurement error in the time dimension (import demand elasticities and AVEs of non-tariff measures are constant) and the fact that over-time variation is very similar across EU countries - product-level imports is the only source of variation across those countries.

[Insert Table 7 here]

[Insert Figure 5 here]

In sum, we find, both across countries and within countries over time, support for the core HO prediction that greater openness to trade tends to improve the earnings of abundant factors relative to scarce factors. That the findings are not stronger is perhaps a result of the crudeness of our openness measure. However, the HO prediction is more clearly supported in these worldwide panel data, particularly by the within-country time-series results, than in most other country-level trade and wages studies (Robbins, 1996; Anderson, 2005; Goldberg and Pavcnik, 2007), possibly because developed countries are over-represented in WIOD, whereas most of the conflicting studies are of developing countries.

C Influence of price-ratio elasticity

The empirical results in Part A and other studies cited there cast serious doubt on the existence of factor price insensitivity and thus on the practical usefulness of the canonical HOS model. However, the evidence in Part B of the effects of openness to trade on relative wages is supportive of the general principles of HO theory. Put together, these two pieces of evidence suggest the usefulness of the GHO model as a framework in which HO theory can be applied to factor prices, much as an essentially similar model proved useful in applying HO theory to the composition of trade in Romalis (2004) and Chor (2010). We can also check the consistency of the data with two testable predictions of the GHO model in section II about variation among countries and over time in the elasticity of relative wages with respect to relative endowments:

1. This elasticity is made smaller (closer to zero) by more openness to trade, because firms have smaller shares of their home market, per-unit trade costs are lower, and trade in intermediates amplifies differences in factor intensity among goods.

2. This elasticity is also made smaller by a higher price-ratio elasticity - the responsiveness of the relative purchaser prices of good to relative factor costs - which permits endowment-absorbing changes in output mix to be achieved with smaller changes in factor prices.

The results in Part B match the first prediction. The positive coefficients on the interactions between openness and endowments indicate not only that greater openness raises the wage of the relatively abundant factor, as predicted by any HO model, but also, as predicted by GHO, that a larger supply of skilled workers lowers the wage of skilled workers by less in an economy that is more open.

It remains to test the second prediction. The price-ratio elasticity depends on (a) the ratio, τ, of the other-cost wedge (OCW) to immobile factor production costs (IFCs) and (b) the degree, η, to which the relative OCWs of goods move with their relative IFCs. If η were zero, the price-ratio elasticity would be approximately $1/(1+\tau)$, which we write for short as $\widetilde{\tau}$.

In the absence of more detailed information on OCWs, our estimation strategy is to measure the size of τ, in aggregate over the whole economy, and to use as a regressor $\widetilde{\tau}$, whose coefficient should pick up (among other things) the average η across the units of observation in the regression. Given this average η, in units with a higher $\widetilde{\tau}$ the endowment-absorbing response of the output mix to changes in endowments should be larger and hence the depressing effect on the relative skilled wage smaller. In our cross-country framework, allowing for both GHO components, we specify the following regression equation:

$$(18)\quad \ln(w_c) = \alpha + \beta_1 \ln(v_c) + \beta_2 \ln(o_c) + \beta_3 (\ln(v_c) \times \ln(o_c)) + \beta_4 \widetilde{\tau}_c + \beta_5 (\ln(v_c) \times \widetilde{\tau}_c) + \varepsilon_c$$

We expect the coefficients (β_3 and β_5) on both interaction terms to be positive, since higher levels of both o_c and $\widetilde{\tau}_c$ offset the negative effects of higher v_c on w_c.

A key issue is evidently how to measure τ at the country level. Simple accounting and the WIOD data tell us what makes up the purchaser price. What requires judgement is the assignment of these elements between OCWs and IFCs.

We can be confident that internal trade costs and taxes belong in OCWs, since they drive a wedge between IFCs and purchaser prices, regardless of the location of the purchaser. WIOD conveniently provides such data in its national Supply-and-Use tables. Specifically, we use the sum of the internal transport margin and net product taxes across sectors in each country and year. On the assumption of this paper that capital is internationally mobile, profits (or, in WIOD, "capital compensation") also belong in OCWs, since they too create a wedge between IFCs and purchaser price regardless of where the product is sold.

Together, these two elements provide a measure of τ which includes OCWs that are mainly 'domestic' ($\tau \equiv \frac{dom.margin+tax+cap.bill}{totalwagebill}$) and relevant to both internal and international transactions. In principle, OCWs should also include foreign trade costs and purchases of traded intermediate goods. However, we have only limited information on foreign trade costs (the international transport margin in WIOD), we cannot satisfactorily identify traded intermediates (many of which are bought domestically rather than imported, especially in large countries), and we cannot measure the amplification of differences in immobile factor intensities as a result of trade in intermediates. Moreover, the effects of foreign trade costs and traded intermediates on the wage-endowment elasticity should be picked up by our openness measure. So we limit our measure of τ to its domestic elements, recognising that it captures only part of the true price-ratio elasticity.

Table 8 reports the cross-country estimates using within-country average values. As earlier, we first estimate our baseline specification with variations (columns (1) to (3)) and then perform robustness tests.

The results are consistent with both GHO predictions. The significant positive coefficient on the interaction term in column (2) shows that a higher price-ratio elasticity (reflecting a lower value of our measure of the ratio of OCWs to IFCs) softens the otherwise inverse relationship between relative wages and relative endowments. In column (3), including the openness variable and its interaction with endowments, the $\widetilde{\tau}$-endowment interaction stays positive and significant at the 10% level, and the openness-endowment interaction is also positive, as GHO would suggest, though less precisely estimated - and similar in magnitude to the specification without $\widetilde{\tau}$ (see Table 6). In columns (4) to (11), we report results of robustness tests. Across all specifications, the coefficients on the endowment-openness interaction are similar to those estimated without including the price-ratio elasticity. The coefficient on the $\widetilde{\tau}$-endowment interaction becomes smaller and less precisely estimated as we add control variables, particularly in the last four columns, because our measure of τ is negatively correlated with GDP per capita (countries with higher OCW/IFC tend to be less developed).[25]

[Insert Table 8 here]

To check that our estimates are not affected by systematic variation in relative wages across sectors, we replicate the cross-section analysis using sector-level wages and controlling for sector-specific effects with sector dummies. Estimates of the baseline specification reported in Table A6 in the Online Appendix confirm the country-level results. The coefficient on the $\widetilde{\tau}$-endowment interaction is slightly lower and less precisely estimated, but still positive as GHO suggests.[26]

[25]If we do not include GDP per capita in the relevant regressions, the coefficient on the $\widetilde{\tau}$-endowment is indeed close to 1.2-1.4 and hence similar in magnitude to the baseline estimate.

[26]Results - available upon request - using the OTRI as a measure of trade barriers are consistent with GHO and convey the same findings as the baseline results.

In Table 9, we report the results of the panel specification. As before, this approach has the advantage of controlling for time-invariant forces by exploiting within-country variation. In Part A, we found no significant association between contemporaneous changes in relative wages and endowments. However, the evidence in Part B suggests that this average effect obscures significant variation with respect to changes in trade openness over time. Here we further extend the empirical setting to allow for changes in $\widetilde{\tau}$.

The estimates in columns (1) and (2) fit the GHO prediction since they show that the wage-endowment elasticity rises significantly as OCW/IFC ratios become lower. Our baseline specification in column (3) shows the effect of interacting endowments and the price-ratio elasticity to be strongly positive, as is that of interacting endowments and openness, both supporting the GHO prediction that higher price-ratio elasticities and greater openness diminish the inverse effect on relative wages of variation in endowments. The panel estimates are more precise than the cross-country ones, though the $\widetilde{\tau}$-interaction term has a lower coefficient - because OCW/IFC ratios vary much more across countries than within them. These estimates can also be used to analyse graphically how the wage-endowment elasticity varies with the two mediating variables: openness and price-ratio elasticity. Since there are two interaction terms, the effect of endowments on wages depends jointly on o and $\widetilde{\tau}$.

Figure 6 shows how the wage-endowment elasticity is affected by the price-ratio elasticity, keeping the level of log openness constant at its median value. The wage-endowment elasticity rises with the price-ratio elasticity, reflecting the interaction effect in column (3). The estimates imply that if, say, Turkey had median level openness and increased its price-ratio elasticity from a low value of 0.3 to 0.5 (i.e. around two standard deviations in the sample), the relative wages of skilled workers would no longer be sensitive to changes in skill abundance.

We then use the same approach to trace out how the wage-endowment elasticity varies with openness. Figure 7 shows the results at the median value of $\widetilde{\tau}$. The wage-endowment elasticity rises faster with openness than with the price-ratio elasticity. As countries reach the average (adjusted) trade/output ratio (around 1), keeping the price-ratio elasticity at its median value, they move into a region of factor price insensitivity, where the skill premium does not correlate with skill abundance. We can use our estimates also to assess factor price sensitivity in an 'almost closed' economy. Keeping the price-ratio elasticity at its median value, at the minimum level of openness in the sample (adjusted trade/output = 0.44) the estimated wage-endowment elasticity is -0.35. In a totally closed economy (e.g. a trade/output ratio of 0.01), the estimates imply that a 10% increase in relative skill abundance would depress the skilled wage premium by a large 16%.

The remaining columns of Table 9 report the results of robustness tests. Controlling for the composition of unskilled labour almost halves the coefficient of the $\widetilde{\tau}$-openness interaction, a results that is however due to the very restricted sample (only 1995, 2000, 2005 and 2009). A similar pattern emerges when we control for GDP per capita and union membership in columns (8) and (9), which is consistent with the weak cross-section evidence. Overall, while the results are in line with both GHO predictions, greater openness has more attenuating effect than a higher price-ratio elasticity on the wage-endowment relationship, though this could be because our measure of $\widetilde{\tau}$ is restricted to its domestic elements.

[Insert Table 9 here]

[Insert Figure 6 here]

[Insert Figure 7 here]

Table A7 in the Online Appendix reports the results of the specification in changes following Blum (2010). There appears to be no mediating role of price-ratio elasticities beyond their contemporaneous effect - the interaction effect is small and usually insignificant in all columns. The openness interaction effect is much the same as in Table A5 (without controlling for $\widetilde{\tau}$): small or no effects with five-year and ten-year windows, but a large positive effect with the 14-year window.

In sum, our evidence supports the predictions of the GHO model concerning the wage-endowment relationship. Both greater trade openness and higher price-ratio elasticities attenuate the negative effect of larger relative skill endowments on the relative wages of skilled workers. GHO predicts factor price sensitivity rather than insensitivity, but high trade openness and low wedges of costs other than immobile factor production costs can cause wages to change only slightly or even not at all with changes in endowments.

D Structural testing

The reduced-form estimates provide support for the GHO model as they show a significant reduction in the negative relationship between relative wages of skilled workers and skill endowments when countries are more open to trade and labor costs represent a higher share of production costs. To scrutinise further the model predictions, we now estimate directly the implied wage-endowment relationship in equation (13), rewritten as follows:

$$(19)\quad \left(\sum_{i=1}^{n}\left[\lambda_{Hj}\left(1-\theta_{Hj}\right)+\lambda_{Lj}\theta_{Hj}\right]\sigma_{j}\right)\left(\widehat{w}_{L}-\widehat{w}_{H}\right)+ \\ +\left(\sum_{j=2}^{n}\left(\lambda_{Hj}-\lambda_{Lj}\right)\epsilon_{j1}\delta_{j1}\left(\theta_{Hj}-\theta_{H1}\right)\right)\left(\widehat{w}_{L}-\widehat{w}_{H}\right)=\widehat{v}_{H}-\widehat{v}_{L}$$

We can refer to the sum on the left-hand side of the equation as the predicted changes in the relative demand for skills as implied by the GHO model. In turn, this term equals the sum of two terms, which have a clear theoretical interpretation as outlined in the theory part. Specifically, the first term in the sum identifies the 'technique' channel through which firms adjust to the initial change in the relative supply of skills by substituting one type of labor for another within each sector j, as mandated by the technology parameter σ_j. This channel absorbs also other *within*-sector adjustments, such as intra-sectoral product mix changes, which have been shown to play a quantitatively important role in HO theory (Schott, 2003). The second term is the 'output mix' adjustment channel, which measures the extent to which relative labour demand adjusts because of the induced changes in the output structure *between* sectors.

We estimate the predicted changes in the relative demand for skills (the left-hand side sum of equation (19)) and see if it correlates with the observed changes in the relative supply of skills. In doing so, we will also separate the predicted changes due to the technique channel from those due to the output-mix one. The objective is to see whether the changes in relative skill endowments as predicted by the GHO model can match the observed changes in endowments.[27]. Given the use of estimated parameters that rely on a number of assumptions and approximations, we take the evidence from this exercise as suggestive (and preliminary) of how the GHO model can match central tendencies in the data.

To operationalise our exercise, we need to measure the different parameters on the left-hand side of equation (19). The factor use and cost shares (the λ's and θ's parameters) can be computed directly from the WIOD Socio-Economic accounts at the sector-level for each country and year in the sample. For each observation, we calculate direct use and compensation for skilled and unskilled labour as a share of total labour use and compensation (the two types of labour being assumed to be the only immobile factors). To measure the price-ratio elasticity δ_{j1}, we make the simplifying assumption that all OCW's are per-unit

[27]This exercise is thus similar in spirit to Blum (2010), who nevertheless decomposes the factor market clearing condition without imposing any economic structure.

and hence set $\eta_{j1} = 0$. The price-ratio elasticity can thus be approximated as $1/(1+\tau_{j1})$, where $\tau_{j1} \equiv \sqrt{\tau_j \tau_1}$[28]. As explained in the previous empirical analysis, we include internal trade costs, (net) taxes and capital compensation in the OCWs.

We are now left with the elasticity of substitution parameters σ's and ϵ's. To estimate the elasticity of substitution in production, σ, we resort to the much used CES aggregator of skilled and unskilled labor. Under this framework, the elasticity can be retrieved from different combinations of the cost minimisation first-order conditions. Here we use the theoretical relationship between relative compensation to skill and relative wage of skill workers, although using other conditions yields similar estimates (available upon request). While most of the existing literature has estimated a single aggregate elasticity between skilled and unskilled labour using both aggregate and subnational data (see e.g. Ciccone and Peri, 2005; and Angrist, 1995), here we are interested in sector-specific elasticities. We thus average the data within each country and sector and estimate sector-specific regressions of relative compensation to skill onto relative wages of unskilled workers at the sector level. Notice that this exercise, while apparently similar to, is fundamentally different from what we are ultimately after, namely the responsiveness of relative wages to changes in labour endowments at the country-level. The underlying and widely accepted assumption in HO-type trade models (see e.g. Leamer and Levinsohn, 1995; and Slaughter, 1997) is that at the national level wages endogenously respond to exogenous changes in factor endowments, while at the industry level relative labour demand responds to changes in relative wages which are regarded as exogenous to that industry. The estimated sector-specific σ's are reported in column (1) of Table A8.

The median elasticity across sectors is 1.59, which is very close to the 1.41 found by Katz and Murphy (1992) for the U.S. and within the range of 1 to 3 that Katz and Autor (1999) consider plausible. Yet, this masks a lot of heterogeneity across sectors, especially between manufacturing and services. The elasticity is significantly higher than 1 in all manufacturing sectors, while the Cobb-Douglas specification implying $\sigma = 1$ cannot be rejected in all service sectors (with the exception of Social and Personal Services), in Agriculture and Mining. In the Finance and Public Administration sectors, the estimates imply an implausible negative value for σ which is however not significantly different from zero. In the empirical analysis, we thus set $\sigma = 0$ for these two sectors.

As for ϵ, we first estimate the two demand parameters γ and β. We apply the estimation approach pioneered by Feenstra (1994) and estimate β's varying across sectors

[28] As shown in Wood (2012), this result follows from assuming that the ratio of immobile factors costs across the two sectors equal the ratio of OCW's

and purchasing country. A variety is thus defined as good j (i.e. a WIOD sector) sold by country z to country $\check{z}$, where z=$\check{z}$ for a domestic variety. The value of total bilateral shipments at the sector level is taken from the WIOD international input-output table. We follow Patel et al. (2014) and proxy prices with the sectoral price deflators from WIOD, averaged across type of use (intermediate and final). The CES demand system is estimated with the LIML estimator and constrained search algorithm introduced by Soderbery (2015) to ensure that $\beta > 1$. The median of the estimated β's across countries for each sector is reported in column (3) of Table A8.

The overall median value is 2.23, which is in line with the estimated elasticities available in the literature[29]. Here there does not seem to be a particular difference between manufacturing and service sectors, although estimates for the service sectors are less reliable due to the relatively small trade and notoriously lower quality of the data - which may also explain the very high median β in the Public Administration sector. The values for the γ's are set to the minimum of the β's for each country and vary from 1.2 for Italy to 7.8 for Bulgaria. These estimated γ's and β's are then combined to form the aggregated elasticity ϵ using the simplified weighed average in equation (11).

Finally, estimation of the predicted changes in relative skill endowments requires choosing a reference sector. With sector-varying elasticities σ's, ϵ and δ's, this choice inevitably affects the values of the technique and output-mix channels. We avoid choosing among service and primary sectors given the few problematic estimated elasticities and the usually low accuracy of trade data in services. The estimates reported below thus rely on choosing Plastic as the reference sector. While choosing other sectors (including services) does affect the point estimates discussed below, the main findings are not affected - results available upon request.

Similarly to the previous empirical analysis, we carry out the structural tests exploiting both cross-country and panel data. In the cross-country analysis, the different components of equation (19) are computed using data averaged within each country-sector over time. Changes in relative wages and skill endowments are computed relative to the average relative wages and endowments across countries. A constant term is added to all regressions so that the choice of the reference country is inconsequential. Further, the predicted changes in the relative demand for skill are arguably measured with error given the combination of many estimated parameters. To control for the influence of outliers in the predicted changes within our small sample of 40 countries, we apply a 'robust' estimator which essentially downweights

[29] Broda and Weinstein (2006), for instance, obtain a median estimated β's (what they refer to as σ's) of 3.39 using a similar methodology on more detailed trade data (excluding hence domestic shipments) for 73 countries.

outliers in fitting the linear regression[30]

We explore the fit of different versions of GHO, where the key element is that elasticities in demand are finite. We start from situation where all the elasticity parameters are set to 1 and per-unit trade costs are zero (i.e. $\sigma_{j1} = \beta_{j1} = \gamma = \delta_{j1} = 1$). We refer to this as the "Cobb-Douglas" (CD) scenario. Under these circumstances, the technique and output-mix terms sum up to 1 and changes in relative skill endowments should be matched one-to-one with changes in the relative wage of unskilled workers. The first scatter plot in the top panel of Figure 8 shows the linear prediction of a robust regression of $(\widehat{w}_L - \widehat{w}_H)$ onto $\widehat{v}_H - \widehat{v}_L$. Not surprisingly, the slope coefficient is lower than 1 and also slightly lower (in absolute value) than the 0.27 reported in column (1) of Table 2 because of the downweighting of outliers - e.g. Brazil is dropped from the regression. The other two plots in the top panel of Figure 8 split the predicted changes in relative skill endowment in the technique and output-mix components. The within-sector term drives most of the observed wage-endowment relationship, with a slope coefficient being very close to the one estimated using total predicted changes. Changes in the relative demand for skills due to adjustment to the output mix are too low to explain the wide variation in relative skill endowments across countries. The slope coefficient is 0.02, but still significantly different from zero. These findings are consistent with the analysis of Blum (2010) despite coming from a different approach.

We next allow substitution to be non-unitary (thus the 'N-US' label) using the estimated σ's on the production side and the β's on the demand side. There is hence no trade protection of the home market ($\epsilon = \beta$) and trade costs are still only ad-valorem or of iceberg type. The performance of the N-US model is better than the one of the previous CD-type scenario (see second panel of Figure 8). The slope between predicted and actual changes in endowments increases to 0.28. Interestingly, the predictive power of the technique channel does not change when we include the sector-varying estimated σ's, perhaps hinting at the fact that for many sectors (including all services) the production technology is not significantly different from Cobb-Douglas. Predicted changes in the relative demand for skills due to adjustment in the output mix across sectors increase on average and now the slope of the linear relationship with actual changes in endowments goes up to 0.15. While overall the model is still far from explaining actual changes in labour endowments across countries, these results suggest that introducing a finite and sector-varying elasticity in the demand for goods and for factors makes output mix adjustments almost as important as adjustment

[30]The estimates are produced using the *rreg* command in STATA - see Verardi and Croux (2009) for details.

through production techniques.

As an intermediate step, we then allow for protection in the home market (or 'Home Bias'), so that the relevant elasticity in demand becomes ϵ. The overall fit of the model slightly improves as shown by increase in the R^2 in the regression of predicted vs. actual changes in skill endowments. Here and in the next test the middle plots showing the relationship between the technique channel and changes in skill endowments do not change since we maintain the sector-specific estimated σ's. Using ϵ's rather than the β's lowers the slope coefficient, although not significantly so.

Finally, we introduce the price-ratio elasticity assuming that all OCW's are of per-unit type (hence the 'P-RE' label). The R^2 of a regression of model-mandated changes in relative skill endowments onto actual changes goes up to 0.42 (left-most plot at the bottom of Figure 8). This slight increase in the fit is mainly related to predicted changes due to output-mix adjustment. As shown by the right-most plot in the bottom panel of Figure 8, the slope relating the output-mix term and actual changes in relative skill endowment goes down to 0.07 when introducing price-ratio elasticities.

Overall, the cross-country slope tests suggest that different versions of GHO can help relating output-related adjustments in relative demand for skill to actual change in its relative supply. Yet, the model explains only partially changes in skill endowments across countries. In the following, we perform similar tests exploit changes in the relative supply of skills within countries over time.

[Insert Figure 8 here]

To this end, we define changes in relative wages and labour endowments as annualised differences over windows of 5, 10 and 14 years, as in our previous empirical analysis and similarly to Blum (2010). The technique and output-mix terms are averaged over the same windows. Table 10 reports the estimated coefficients of regressions of predicted changes in the relative demand for skills from the model, the technique term and the output-mix one onto observed changes in relative skill endowments. The 5- and 10-year regressions control for country-specific linear trends and year dummies - see also equation (16).

Tests of the model with unitary elasticities (Cobb-Douglas columns) confirms the cross-country evidence. The technique term explains most of the relationship between predicted and actual changes in skill endowments, which, as already shown in Table 4, is less than one and drops to zero with the longest differences (14 years). Changes due to adjustments in the output mix do however predict observed changes in skill endowments, with the slope being slightly higher than in the cross-section.

When we introduce non-unitary elasticities (N-US columns), the estimated slope goes up substantially in both 5- and 10-year specifications, while remaining below one and being imprecisely estimated. As in the cross-section, this increase is driven by the output-mix term, since allowing for sector-varying elasticities in production does not affect the predictive power of the technique term.[31] Yet, the arguably high noise in the estimated elasticities inflates standard errors and lowers the overall fit of the model, as shown by the Within R^2.

Moving to a specification with trade protection (Home Bias columns), predicted changes in the relative demand for skills are smoothed out and their relationship with actual changes in endowments is more precisely estimated. The slope coefficient is still high and adjustments in the output-mix becomes at least as important as within-sector adjustments with 5- and 10-year changes. The model however has no predictive power as we go to 14-year changes.

Finally, results are qualitatively confirmed when we adopt a full GHO specification with per-unit OCW's (P-RE columns). The slope is slightly lower (although not in a statistical sense) than with only ad-valorem trade costs, but the fit of the model (and of its components) remains the same.

[Insert Table 10 here]

While subject to a number of caveats and still preliminary, these results suggest that our proposed version of the GHO, when taken at 'face value', can explain a non-trivial amount of the observed variation in relative skill abundance across countries and, more importantly, over time. While most of this variation remains unexplained, finite and heterogeneous demand elasticities seems to account, at least partly, for adjustments in the labour market within an open economy. While allowing for a less-than-unitary price-ratio elasticity does not explain adjustments in the relative demand for skills, softening the stark assumption of zero-or-all per-unit OCW's (i.e. allowing $\eta \in (0, 1)$) may bring the model closer to reality.

V Concluding remarks

This paper has set out and extended the more general HO (GHO) model that has emerged from recent analytical work on trade, and has applied it to the determination of factor prices in open economies, using WIOD's global panel dataset. The results show that the basic HO prediction that greater openness tends to raise the relative price of a country's abundant

[31] In the N-US specification of the model, the point estimate of the slope varies also more depending on the choice of the numeraire sector.

factor is consistent with both cross-country evidence (rarely used in earlier studies) and country-level time series evidence (more clearly so than in most other recent studies).

However, the GHO model provides a more convincing explanation than the standard HOS model of the observed relationships among wages, endowments and openness, above all because HOS cannot explain (while GHO can) the apparently continuous inverse responsiveness of relative wages to relative skill supplies over time within countries that are open to trade. The results are also consistent with the GHO model's prediction that the impact of endowments on factor prices increases with the height of barriers to trade and decreases with the share of wages in the cost of production.

The relative accuracy of the HOS and GHO models as explanations of reality is of more than academic significance. If the effects on wages and other factor prices of more or less openness to trade are just one element of a broader demand and supply system, as in the GHO model, rather than operating through a rigid link with world prices and trade barriers, as in HOS, the implications and options for policy are different. For example, in GHO a trade-induced widening of the wage gap between skilled and unskilled workers could be reversed by educating and training unskilled workers, which it could not in one-cone HOS. The GHO model could thus enable more constructive dialogue among trade economists, labour economists and policy makers.

As ever, there is much scope for further work. The analysis in this paper could be extended by testing other likely influences on factor prices in the GHO model, including the composition of domestic demand and variations in technology. It could also be complemented by more empirical analysis of relationships between the composition of endowments, output and trade, including further comparisons of the performance of the HOS and GHO models. Another important step would be to use the WIOD data to test systematically the present paper's assumption that capital is an internationally mobile factor, which is at variance with the assumption of most HO empirical studies. More challenging in terms of the availability of data, especially on factor prices, would be to introduce land as another immobile factor.

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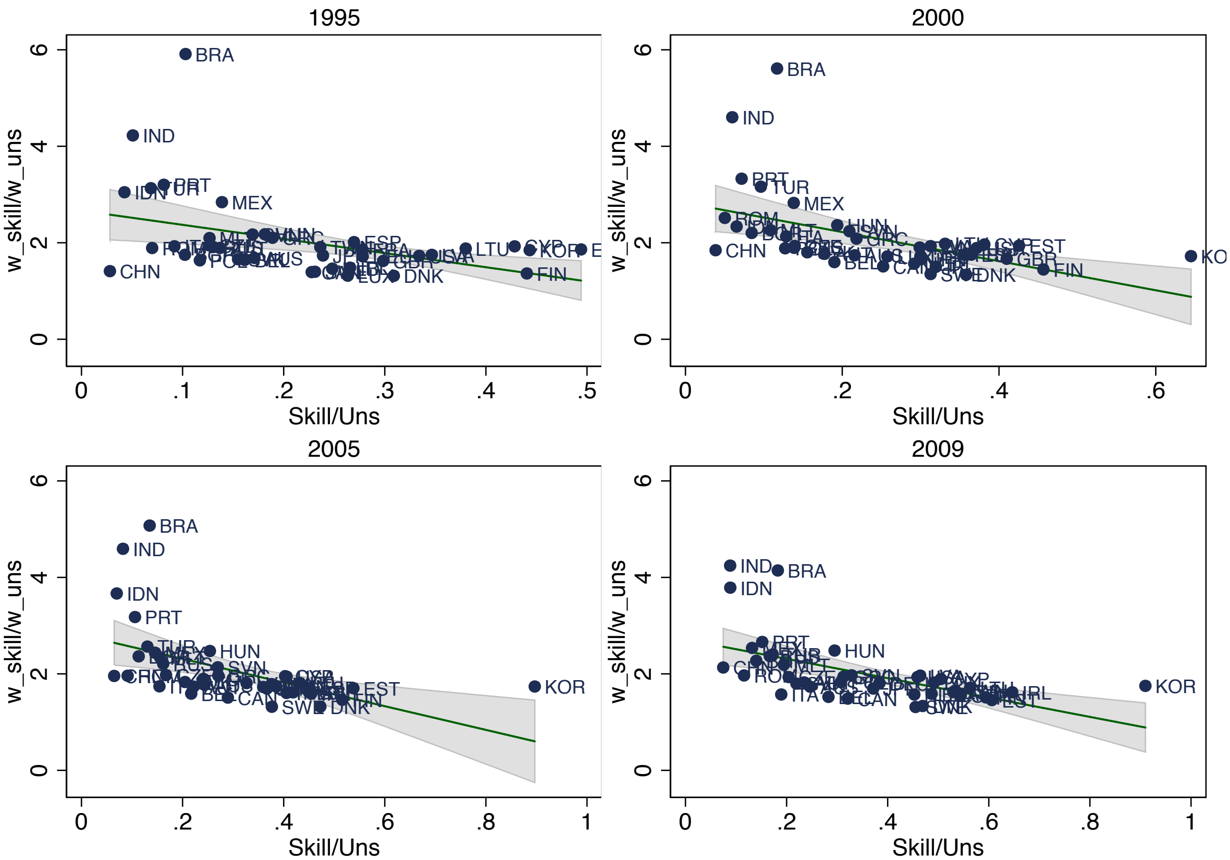
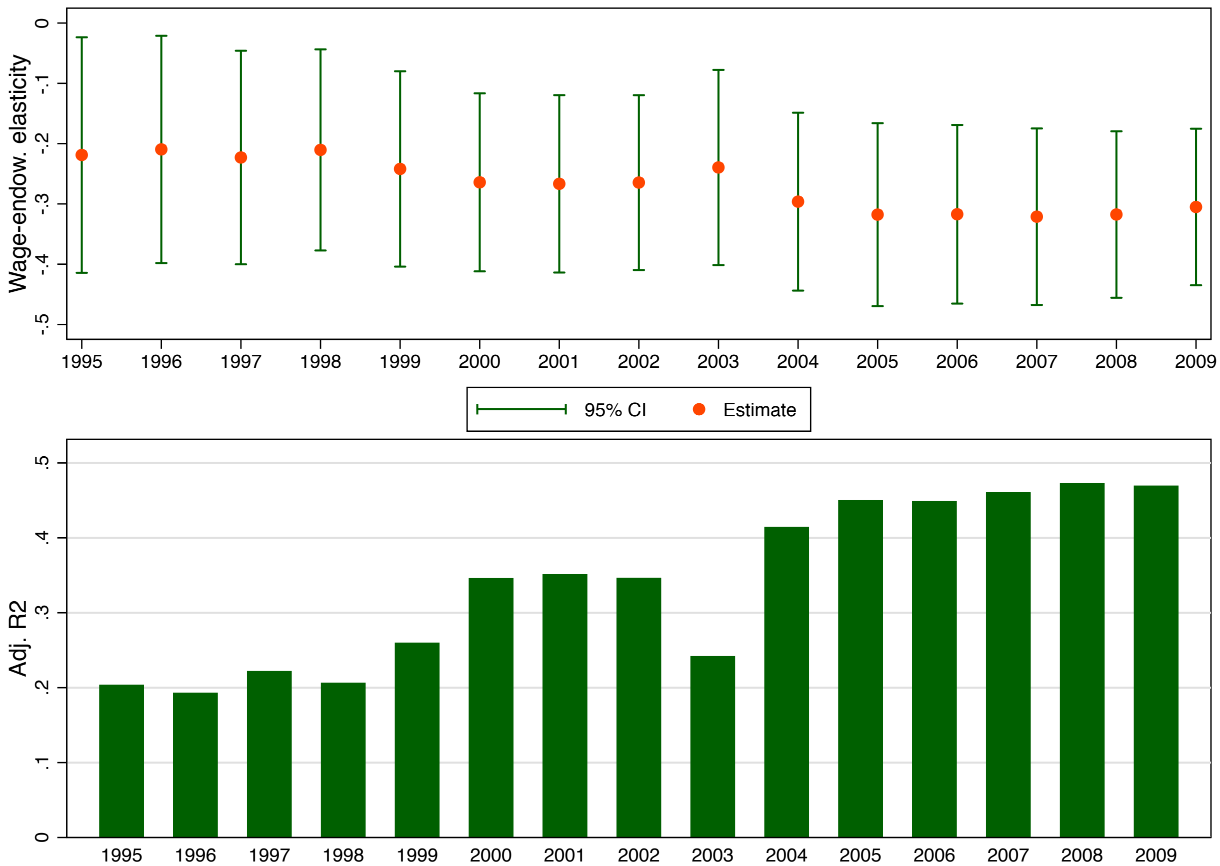
List of figures

Figure 1: Relative wages and endowments



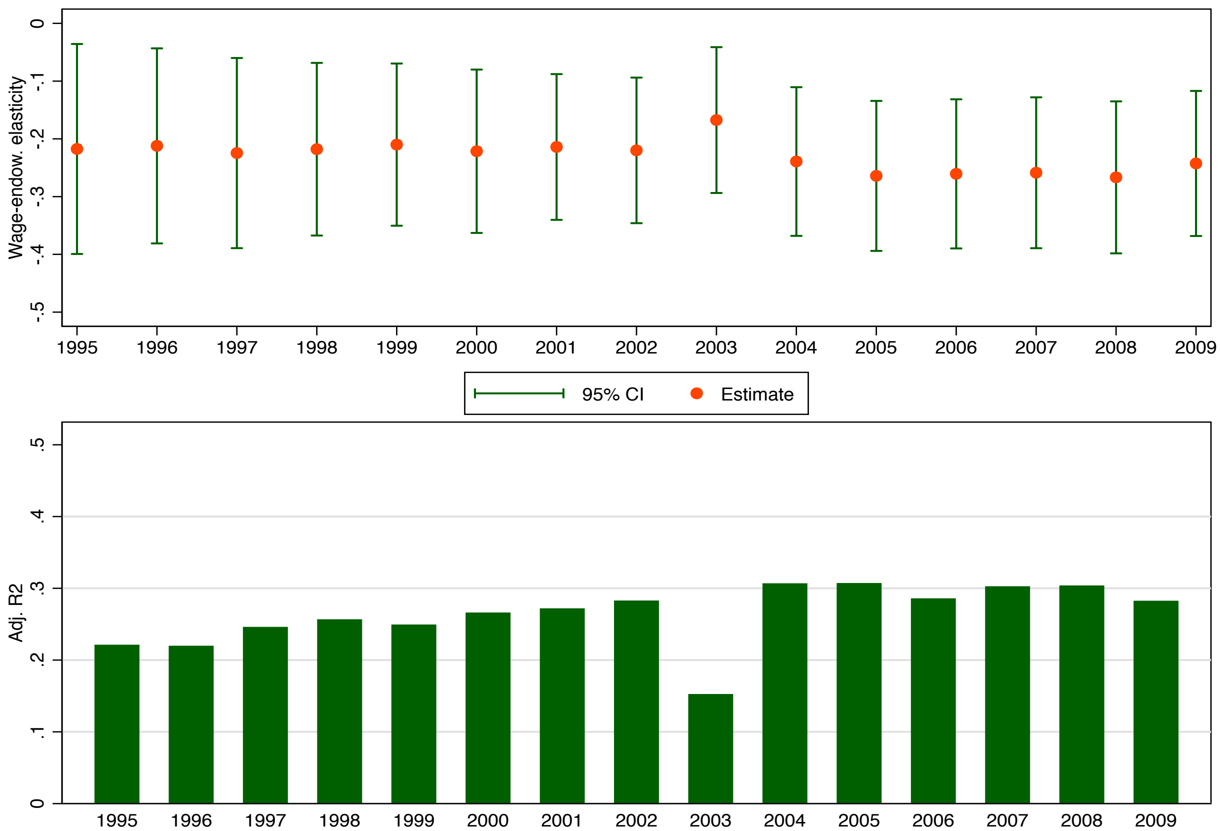
Linear prediction line with 90% confidence bands (based on heteroskedasticity-robust standard errors) are shown.

Figure 2: Wage-endowment regressions - Elasticities and R^2



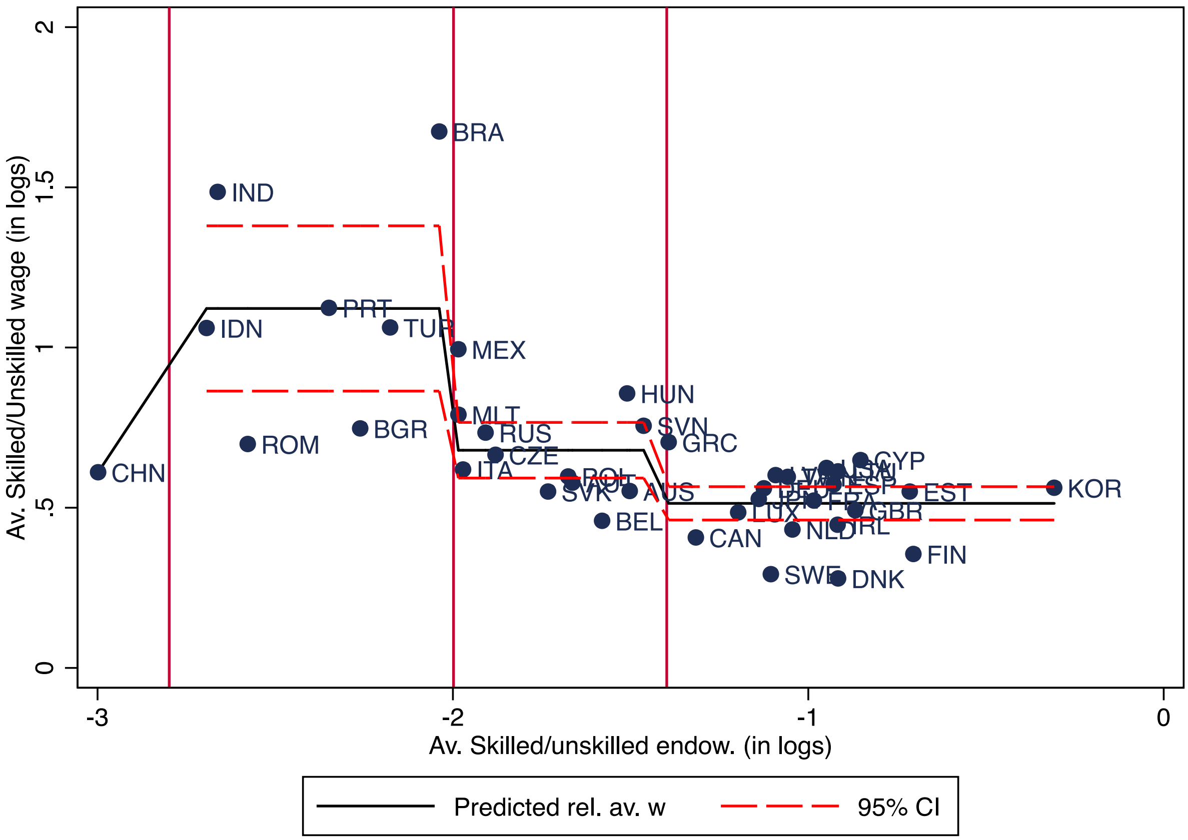
All regressions include a constant term. Confidence intervals are calculated using heteroskedasticity-robust standard errors.

Figure 3: Wage-endowment regressions - Elasticities and R^2 (sectoral wages)



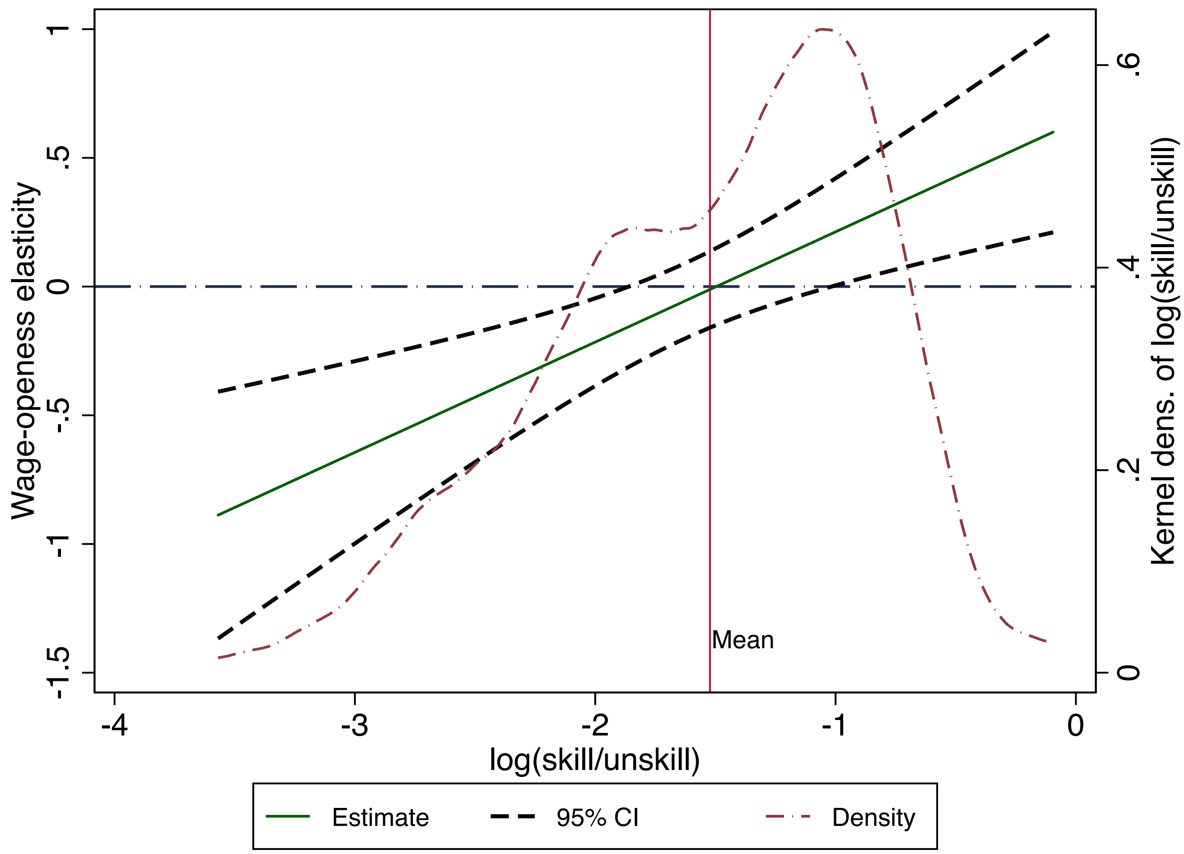
All regressions include sector dummies. Confidence intervals are calculated using standard errors clustered at the country level.

Figure 4: Estimated cones of diversification



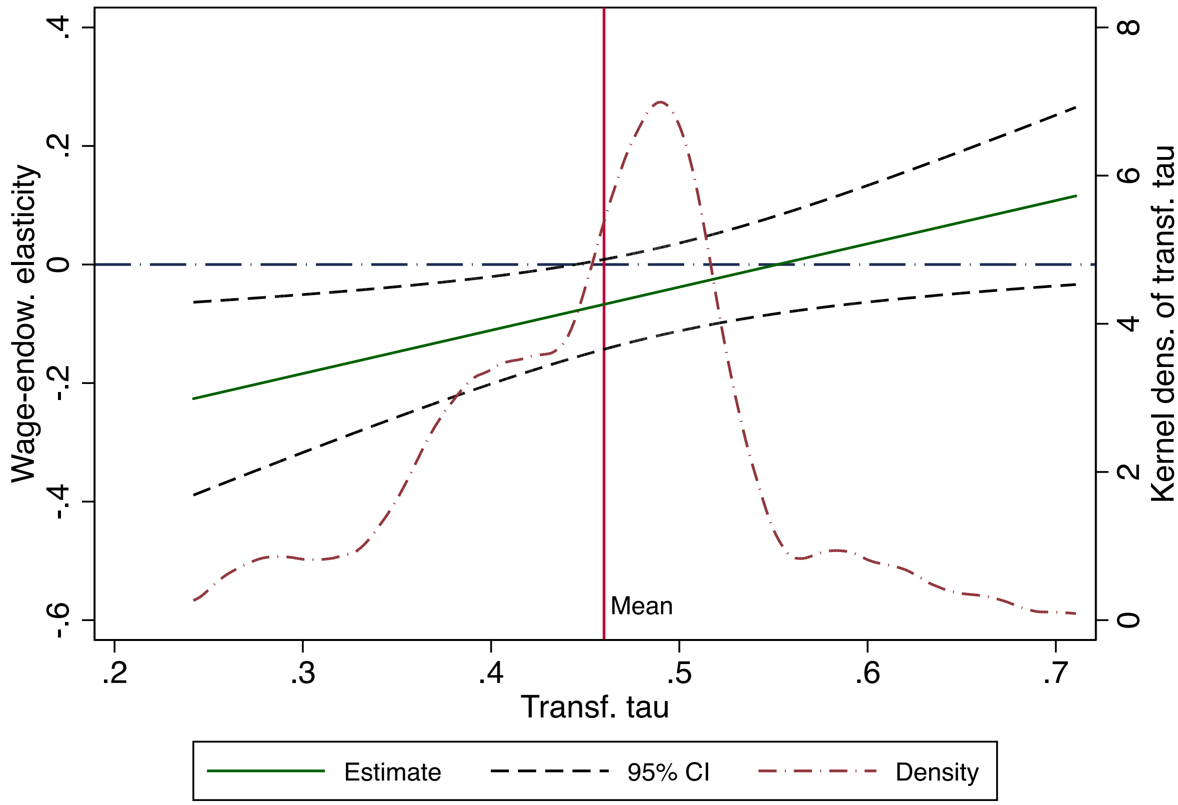
Vertical lines indicate the location of interior knots.

Figure 5: Wage-openness elasticity and endowments



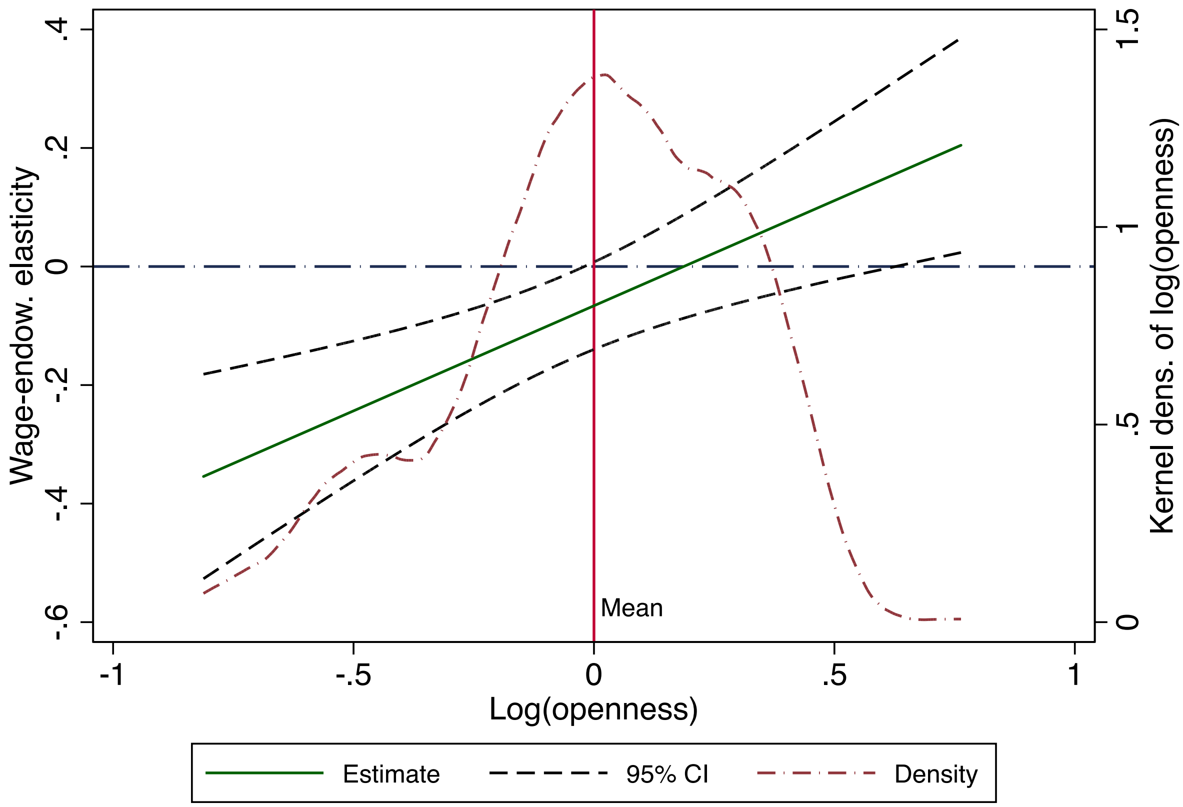
Vertical line is at the sample mean of $\ln(v)$.

Figure 6: Wage-endowment elasticity and $\widetilde{\tau}$ at median openness



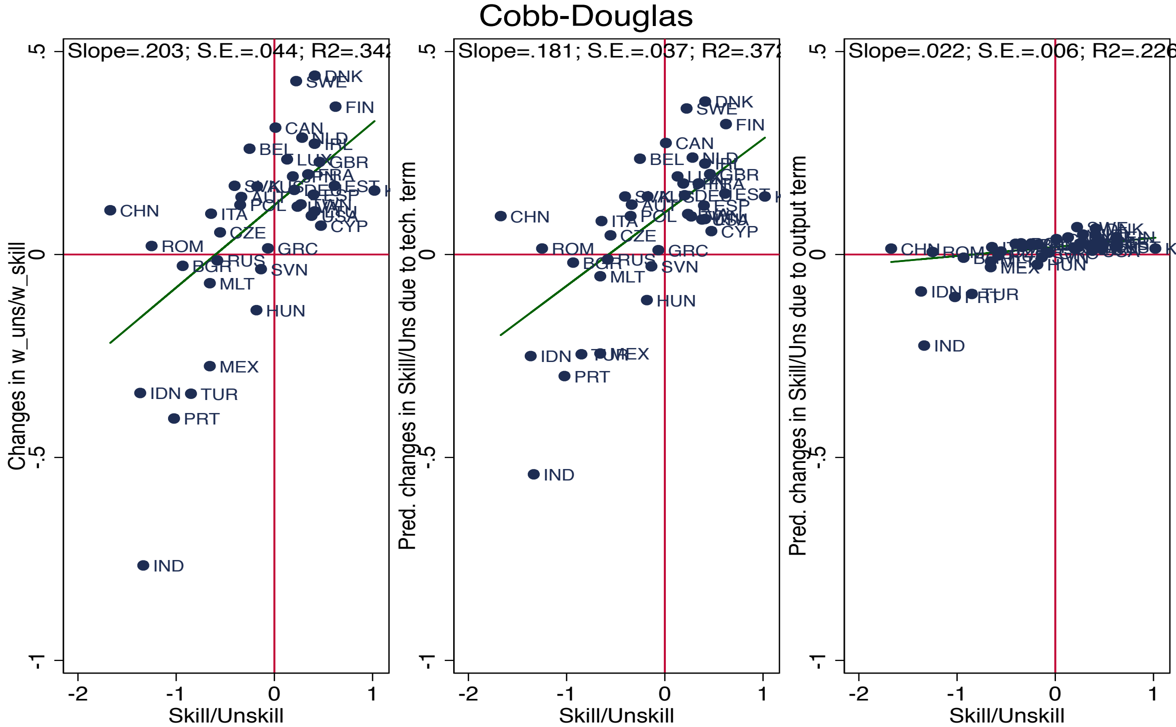
Vertical line is at the sample mean of $\widetilde{\tau}$.

Figure 7: Wage-endowment elasticity and openness at median $\widetilde{\tau}$

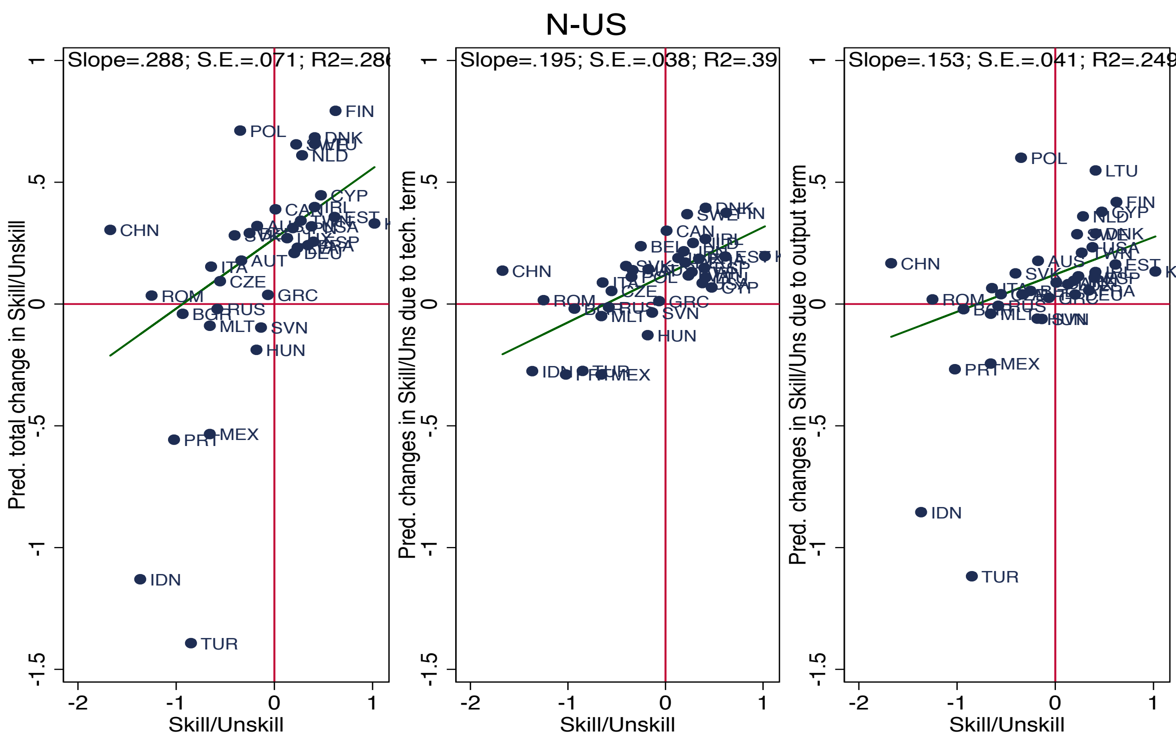


Vertical line is at the sample mean of $\ln(o)$.

Figure 8: Slope tests - Cross-country estimates

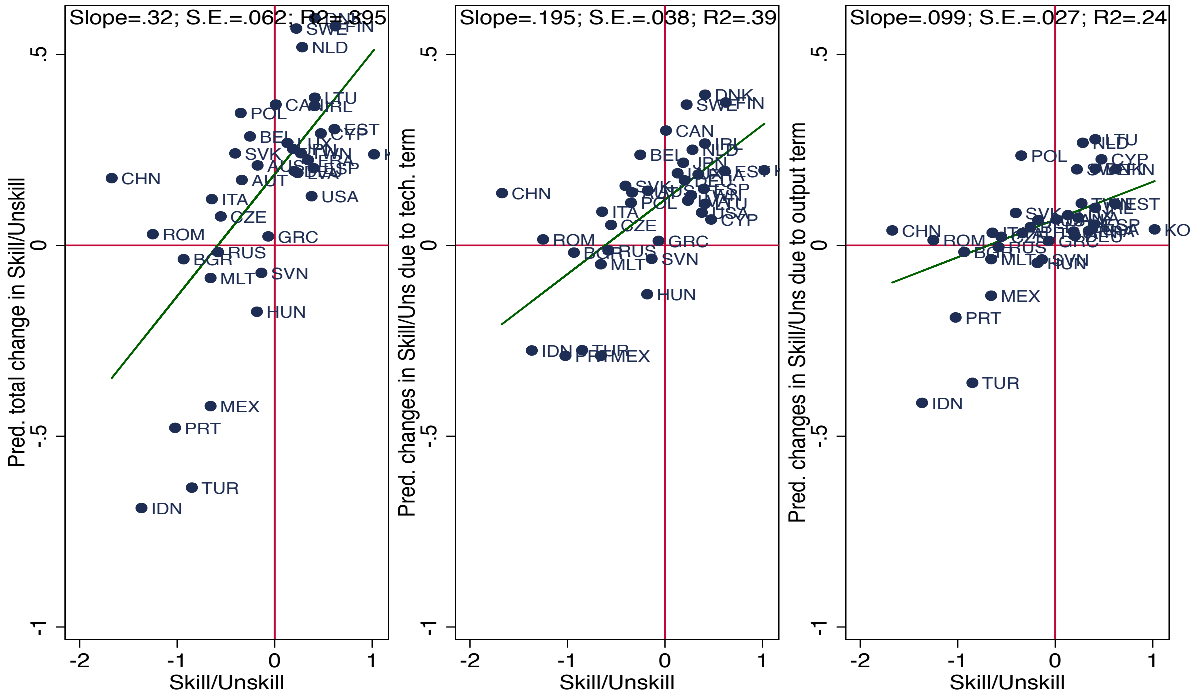


Brazil is excluded since outlier.



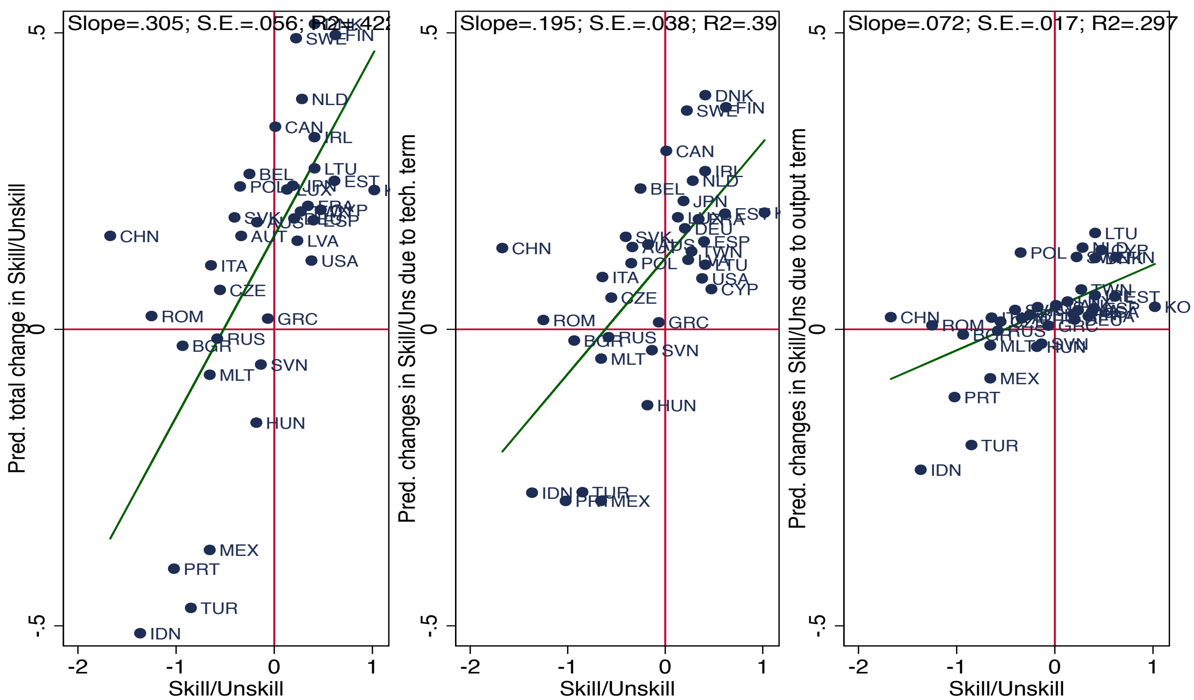
Brazil, UK and India are excluded since outliers.

Home bias



Brazil, UK and India are excluded since outliers.

P-RE



Brazil, UK and India are excluded since outliers.

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Table 1: Summary statistics (country-level)

	Obs	Mean	Std. dev.	Min	Max
$\ln(w)$	600	0.669	0.292	0.244	1.801
$\ln(v)$	600	-1.522	0.665	-3.569	-0.094
$\ln(o)$	600	0.000	0.288	-0.812	0.763
$\widetilde{\tau}$	600	0.457	0.078	0.242	0.711
$\ln(w_m)$	600	0.503	0.267	-0.072	1.509
$\ln(v_m)$	600	0.871	1.241	-1.798	3.702
$\ln(v_{ls})$	600	-1.145	0.774	-3.448	-0.106
$\ln(v_{ns})$	158	-3.736	1.304	-6.738	-0.820
$\ln(v_{land})$	600	0.099	1.290	-3.353	3.600
$\ln(r)$	600	-2.057	0.338	-2.906	-0.891
$\ln(GDPpc)$	600	9.531	1.099	6.152	11.382
$\ln(labmkt)$	200	4.130	0.231	3.658	4.605
$\ln(union)$	434	-1.275	0.638	-3.005	-0.085
$\ln(o_{fin})$	600	0.000	0.280	-0.900	0.556
$\ln(o_{inp}$	600	0.000	0.326	-0.858	0.899
$\ln(OTRI)$	455	-2.261	0.461	-4.006	-0.987

w_m and v_m are relative wages and endowments (respectively), calculated defining 'skilled' labor as the sum of high- and medium-skilled labor.

Table 2: Wage-endowment elasticity - Robustness checks (cross-country)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Skill composition				Other factors			Control variables	
$\ln(v)$	-0.277***	-0.243***	-0.212***	-0.0862***	-0.275***	-0.251***	-0.247***	-0.157**	-0.178***
	(0.0751)	(0.0700)	(0.0601)	(0.0280)	(0.0765)	(0.0719)	(0.0748)	(0.0631)	(0.0635)
$\ln(v_{ls})$		0.0712**	-0.00256						
		(0.0302)	(0.0365)						
$\ln(v_{ns})$			0.0808**						
			(0.0363)						
$\ln(v_{land})$					-0.00744		-0.0102		
					(0.0241)		(0.0250)		
$\ln(r)$						0.107	0.113		
						(0.101)	(0.107)		
$\ln(GDPpc)$								-0.0844	-0.103*
								(0.0588)	(0.0592)
$\ln(labmkt)$								-0.211*	
								(0.115)	
$\ln(union)$									-0.109*
									(0.0628)
Obs	40	40	40	40	40	40	40	40	36
R^2	0.383	0.413	0.471	0.174	0.384	0.394	0.396	0.509	0.527

All variables are within-country averages. Columns (8) and (9) include a dummy for $EU15$ membership. Medium-skilled workers are aggregated to high-skill workers in computing the dependent variable w and variable v in column (4). All regressions include a constant term. Heteroskedasticity-robust standard errors are reported in parenthesis. Significant at: *10%, **5%, ***1% level.

Table 3: Wage-endowment elasticity - Panel estimates

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Skill composition				Other factors		Control vars
$\ln(v)$	-0.0781	-0.0745	-0.112	-0.0685	-0.0756	-0.0717	-0.0506
	(0.0976)	(0.0987)	(0.108)	(0.0825)	(0.0878)	(0.0893)	(0.0775)
$\ln(v_{ls})$		0.0403	-0.0294				
		(0.0942)	(0.0984)				
$\ln(v_{ns})$			0.00491				
			(0.0120)				
$\ln(v_{land})$						-0.0272	
						(0.0780)	
$\ln(r)$					0.197***	0.197***	
					(0.0677)	(0.0671)	
$\ln(GDPpc)$							0.235*
							(0.118)
$\ln(union)$							0.0902
							(0.0558)
Obs	600	600	158	600	600	600	434
Within R^2	0.069	0.071	0.069	0.209	0.141	0.141	0.201

Medium-skilled workers are aggregated to high-skill workers in computing the dependent variable w and variable v in column (4). All regressions include year dummies. Column (7) include a dummy for *EU* membership. Standard errors clustered at the country level are reported in parenthesis. Significant at: *10%, **5%, ***1% level.

Table 4: Wage-endowment elasticity - Panel estimates (in changes)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)
	5-year changes					10-year changes					14-year changes				
$\widehat{v}$	-0.191**	-0.194**	-0.0820	-0.199**	-0.0935*	-0.323***	-0.335***	-0.184*	-0.312***	-0.250***	-0.0126	-0.0134	-0.0276	-0.0278	-0.267**
	(0.0750)	(0.0770)	(0.0825)	(0.0766)	(0.0470)	(0.0926)	(0.104)	(0.0917)	(0.0842)	(0.0561)	(0.144)	(0.147)	(0.109)	(0.135)	(0.123)
$\widehat{v}_{ls}$		-0.0255					-0.0412					-0.0308			
		(0.0564)					(0.0646)					(0.103)			
$\widehat{v}_{land}$				0.0871					0.103					-0.119	
				(0.0839)					(0.110)					(0.114)	
$\widehat{r}$				0.0144					0.0772					0.218**	
				(0.0595)					(0.0660)					(0.0868)	
$\widehat{GDPpc}$					-0.160					-0.0912					0.471***
					(0.201)					(0.122)					(0.0902)
$\widehat{union}$					0.00413					-0.0393					0.262***
					(0.0845)					(0.0504)					(0.0797)
Obs	400	400	400	400	258	200	200	200	200	129	40	40	40	40	24
Within R^2	0.216	0.217	0.298	0.226	0.311	0.313	0.314	0.234	0.344	0.369	0.000	0.002	0.002	0.185	0.521

Medium-skilled workers are aggregated to high-skill workers In computing the dependent variable $\widehat{w}$ and variable $\widehat{v}$ in columns (3), (8) and (14). Columns (1) to (11) include country and year dummies. Standard errors clustered at the country level are reported in parenthesis. Significant at: *10%, **5%, ****1% level.

Table 5: Wage-endowment elasticity - Panel estimates (sector-level wages)

	(1)	(2)	(3)	(4)
	Levels	5-year ch.	10-year ch.	14-year ch.
v	0.0451	0.0391	-0.147*	0.0963
	(0.0548)	(0.0484)	(0.0810)	(0.0936)
Obs	19,753	13,162	6,577	1,314
R^2	0.693	0.235	0.393	0.101

All regressions in columns (1) to (3) include, country, sector and year fixed effects. Regression in column (4) include sector fixed effects. Standard errors clustered at the country-level are reported in parenthesis. Significant at: *10%, **5%, ***1% level.

Table 6: Endowment-openness estimates - Cross-section

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	Baseline			Skill composition		Other factors		Control variables			
$\ln(o)$	-0.179	-0.270	0.480	0.249	0.0929	0.458	0.177	0.711	1.197	0.898*	-1.610
	(0.188)	(0.164)	(0.449)	(0.423)	(0.591)	(0.459)	(1.274)	(0.436)	(2.747)	(0.517)	(3.467)
$\ln(v)$		-0.293***	-0.319***	-0.265***	-0.294***	-0.282***	-0.282***	-0.142**	-0.149**	-0.130**	-0.192***
		(0.0731)	(0.0716)	(0.0581)	(0.0614)	(0.0753)	(0.0791)	(0.0600)	(0.0678)	(0.0627)	(0.0732)
$\ln(o) \times \ln(v)$			0.530	0.420	0.518	0.534	0.504	0.725**	0.606	0.855**	0.571
			(0.374)	(0.346)	(0.319)	(0.377)	(0.454)	(0.347)	(0.408)	(0.396)	(0.441)
$\ln(v_{ls})$				-0.0646	-0.127*						
				(0.0573)	(0.0771)						
$\ln(v_{ns})$				0.103**	0.114**						
				(0.0480)	(0.0550)						
$\ln(v_{land})$						-0.0241	-0.0252				
						(0.0266)	(0.0344)				
$\ln(r)$						0.132	0.137				
						(0.100)	(0.118)				
$\ln(GDPpc)$								-0.157**	-0.171**	-0.193***	-0.189**
								(0.0626)	(0.0807)	(0.0714)	(0.0839)
$\ln(labmkt)$								-0.104	-0.127		
								(0.117)	(0.124)		
$\ln(union)$										-0.0270	-0.0742
										(0.0611)	(0.0573)
Other $\ln(o)$ interactions?	N	N	N	N	Y	N	Y	N	Y	N	Y
Obs	40	40	40	40	40	40	40	40	40	36	36
R^2	0.00515	0.421	0.478	0.553	0.573	0.476	0.444	0.649	0.628	0.694	0.707

All variables are within-country averages. Columns (8) to (11) include a dummy for *EU*15 membership; columns (9) and (11) include also its interaction with $\ln(o)$. All regressions include a constant term. Standard errors bootstrapped with 500 replications are in parenthesis. Significant at: *10%, **5%, ***1% level.

Table 7: Endowment-openness estimates - Panel

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Baseline			Skill comp.		Other factors		Control vars	
$\ln(o)$	-0.134	-0.119	0.641***	0.617***	0.592	0.564***	0.796*	0.786***	-0.288
	(0.109)	(0.105)	(0.199)	(0.195)	(0.443)	(0.191)	(0.427)	(0.174)	(1.531)
$\ln(v)$		-0.0573	-0.103	-0.181**	-0.178**	-0.0891	-0.0910	-0.0758	-0.0765
		(0.0947)	(0.0848)	(0.0853)	(0.0888)	(0.0828)	(0.0789)	(0.0535)	(0.0538)
$\ln(o) \times \ln(v)$			0.428***	0.438***	0.437***	0.390***	0.400***	0.474***	0.303*
			(0.113)	(0.110)	(0.148)	(0.111)	(0.104)	(0.0982)	(0.174)
$\ln(v_{ls})$				-0.0388	-0.0326				
				(0.0771)	(0.0898)				
$\ln(v_{ns})$				0.00316	0.00208				
				(0.0104)	(0.0130)				
$\ln(v_{land})$						-0.0441	-0.0402		
						(0.0719)	(0.0712)		
$\ln(r)$						0.155**	0.152**		
						(0.0654)	(0.0663)		
$\ln(GDPpc)$								0.164	0.180*
								(0.102)	(0.0965)
$\ln(union)$								0.0853*	0.117**
								(0.0471)	(0.0523)
Other $\ln(o)$ interactions?	N	N	N	N	Y	N	Y	N	Y
Obs	600	600	600	158	158	600	600	434	434
R^2	0.923	0.923	0.935	0.936	0.935	0.939	0.939	0.946	0.948

All columns include year dummies. Columns (8) and (9) include a dummy for *EU* membership; column (9) includes also its interaction with $\ln(o)$. Standard errors bootstrapped with 500 replications and country-level clustering are in parenthesis. Significant at: *10%, **5%, ***1% level.

Table 8: Wage-endowment elasticity, price-ratio elasticity and openness - Cross-section

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	Baseline			Skill comp.		Other factors		Controls vars			
$\ln(v)$	-0.241***	-1.028***	-1.037**	-0.757*	-1.100**	-0.967**	-1.038*	-0.659	-1.664	-0.464	-0.914
	(0.0760)	(0.259)	(0.415)	(0.391)	(0.500)	(0.427)	(0.561)	(0.424)	(1.402)	(0.423)	(1.087)
$\ln(\widetilde{\tau})$	-0.708	1.668***	1.624	0.908	1.229	1.384	1.412	1.456	0.867	1.080	0.330
	(0.537)	(0.590)	(1.118)	(1.077)	(1.370)	(1.194)	(1.937)	(1.409)	(2.301)	(1.514)	(2.628)
$\ln(v) \times \ln(\widetilde{\tau})$		1.718***	1.647*	1.126	1.209	1.531*	1.551	1.089	0.944	0.704	0.106
		(0.555)	(0.874)	(0.836)	(0.963)	(0.900)	(1.407)	(0.932)	(1.402)	(0.945)	(1.750)
$\ln(o)$			0.389	0.221	0.488	0.336	0.401	0.639	0.714	0.862	1.119*
			(0.389)	(0.410)	(0.512)	(0.450)	(0.543)	(0.455)	(0.585)	(0.526)	(0.647)
$\ln(v) \times \ln(o)$			0.469	0.390	0.550	0.454	0.498	0.674*	0.709	0.833**	0.980*
			(0.327)	(0.341)	(0.393)	(0.375)	(0.443)	(0.366)	(0.487)	(0.397)	(0.500)
$\ln(v_{ls})$				-0.0431	0.0217						
				(0.0620)	(0.177)						
$\ln(v_{ns})$				0.0755*	-0.0839						
				(0.0444)	(0.102)						
$\ln(v_{land})$						-0.0233	0.00135				
						(0.0233)	(0.109)				
$\ln(r)$						0.0282	-0.0263				
						(0.123)	(0.419)				
$\ln(GDPpc)$								-0.142*	-0.00953	-0.188**	-0.0155
								(0.0735)	(0.185)	(0.0736)	(0.201)
$\ln(labmkt)$								-0.106	0.0568		
								(0.127)	(0.418)		
$\ln(union)$										-0.0153	0.161
										(0.0618)	(0.175)
Other $\ln(v)$ interactions?	N	N	N	N	Y	N	Y	N	Y	N	Y
Obs	40	40	40	40	40	40	40	40	40	36	36
R^2	0.379	0.457	0.567	0.591	0.628	0.551	0.524	0.666	0.659	0.689	0.712

All variables are within-country averages. Columns (8) to (11) include a dummy for $EU15$ membership; columns (9) and (11) include also its interaction with $\ln(v)$. All regressions include a constant term. Standard errors bootstrapped with 500 replications are in parenthesis. Significant at: *10%, **5%, ***1% level.

Table 9: Wage-endowment elasticity, price-ratio elasticity and openness - Panel

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Baseline			Skill comp.		Other factors		Controls vars	
$\ln(v)$	-0.0698	-0.484***	-0.411***	-0.300*	-0.215	-0.439***	-0.377***	-0.314*	-0.245
	(0.0488)	(0.170)	(0.157)	(0.169)	(0.157)	(0.135)	(0.135)	(0.163)	(0.431)
$\ln(\widetilde{\tau})$	-1.004***	0.636	0.461	0.0145	0.0352	0.903*	1.627*	-0.00267	0.100
	(0.286)	(0.601)	(0.541)	(0.612)	(0.619)	(0.548)	(0.876)	(0.530)	(0.583)
$\ln(v) \times \ln(\widetilde{\tau})$		0.881**	0.730**	0.379	0.356	0.771***	1.170**	0.399	0.398
		(0.339)	(0.312)	(0.341)	(0.349)	(0.259)	(0.499)	(0.290)	(0.323)
$\ln(o)$			0.490***	0.448**	0.382*	0.502***	0.551***	0.724***	0.612***
			(0.182)	(0.175)	(0.204)	(0.171)	(0.169)	(0.175)	(0.169)
$\ln(v) \times \ln(o)$			0.355***	0.352***	0.318***	0.355***	0.382***	0.452***	0.392***
			(0.103)	(0.104)	(0.112)	(0.104)	(0.102)	(0.0959)	(0.102)
$\ln(v_{ls})$				-0.119	-0.207**				
				(0.0767)	(0.0978)				
$\ln(v_{ns})$				0.00104	0.0533*				
				(0.0122)	(0.0310)				
$\ln(v_{land})$						-0.0593	-0.0283		
						(0.0696)	(0.0964)		
$\ln(r)$						0.117*	0.302*		
						(0.0695)	(0.171)		
$\ln(GDPpc)$								0.0877	0.127
								(0.0663)	(0.0961)
$\ln(union)$								0.0479	0.148
								(0.0503)	(0.0963)
Other $\ln(v)$ interactions?	N	N	N	N	Y	N	Y	N	Y
Obs	600	600	600	158	158	600	600	434	434
R^2	0.924	0.928	0.938	0.936	0.938	0.940	0.940	0.944	0.948

All columns include year dummies. Columns (8) and (9) include a dummy for *EU* membership; column (9) includes also its interaction with $\ln(v)$. Standard errors bootstrapped with 500 replications and country-level clustering are in parenthesis. Significant at: *10%, **5%, ***1% level.

Table 10: Slope tests - Panel estimates

Dep. var.: Pred. $\widehat{v}$	Indep. var.: Observed $\widehat{v}$												
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
	Cobb-Douglas			N-US			Home bias			P-RE			
	Slope	SE	R^2	Slope	SE	R^2	Slope	SE	R^2	Slope	SE	R^2	Obs
5-year changes													
Total	0.191*	(0.0750)	0.216	0.356	(0.256)	0.138	0.344*	(0.152)	0.191	0.286*	(0.119)	0.206	400
Tech.	0.171*	(0.0691)	0.215	0.178*	(0.0736)	0.205	0.178*	(0.0736)	0.205	0.178*	(0.0736)	0.205	400
Output	0.0575*	(0.0226)	0.200	0.178	(0.191)	0.100	0.166	(0.0822)	0.147	0.108*	(0.0480)	0.160	400
10-year changes													
Total	0.323**	(0.0926)	0.313	0.799*	(0.373)	0.168	0.619**	(0.214)	0.267	0.478**	(0.154)	0.287	200
Tech.	0.265***	(0.0738)	0.316	0.282**	(0.0826)	0.297	0.282**	(0.0826)	0.297	0.282**	(0.0826)	0.297	200
Output	0.0900**	(0.0293)	0.279	0.516	(0.303)	0.116	0.337*	(0.140)	0.207	0.195*	(0.0791)	0.213	200
14-year changes													
Total	0.113	(0.115)	-0.000668	0.0704	(0.266)	-0.0244	0.0514	(0.186)	-0.0243	0.0803	(0.150)	-0.0186	40
Tech.	0.0838	(0.0918)	-0.00426	0.131	(0.0979)	0.0199	0.131	(0.0979)	0.0199	0.131	(0.0979)	0.0199	40
Output	0.00834	(0.0213)	-0.0222	0.0403	(0.122)	-0.0234	-0.0139	(0.0771)	-0.0254	-0.00375	(0.0440)	-0.0261	40

All regressions in 5- and 10-year changes include country and year dummies. Standard errors are clustered at the country level. R^2 are within countries. Estimates of the specification in 14-year changes are from robust regressions. Adjusted R^2 is reported. Significant at: *10%, **5%, ***1% level.