Essays in international economics and macroeconomics

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ABSTRACT

ESSAYS IN INTERNATIONAL ECONOMICS AND MACROECONOMICS

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The present dissertation is composed by three essays. The first essay is titled **"Comparative Advantage, Service Trade, and Global Imbalances"**.

The large current account deficit of the U.S. is the result of a large deficit in the goods balance and a modest surplus in the service balance. The opposite is true for Japan, Germany and China. Moreover, I document the emergence from the mid-nineties of a strong negative relation between specialization in export of services and current account balances in a large sample of OECD and developing countries.

Starting from these new stylized facts, I propose in this essay a "service hypothesis" for global imbalances, a new explanation based on the interplay between the U.S. comparative advantage in services and the asymmetric trade liberalization process in goods trade versus service trade that took place in the last 15 years. I use a structural gravity model to quantify the extent of this asymmetry. I show that a simple two-period model can rationalize the emergence of current account deficits in the presence of such asymmetric liberalization. The key inter-temporal mechanism is the asymmetric timing of trade policies, which affects savings decisions.

Finally, I explore the quantitative relevance of this explanation for global imbalances. A multi-period version of the model, fed with the asymmetric trade liberalization path found

in the data, generates a current account deficit of about 1% of GDP (roughly 20% of what was observed in the U.S. in 2006).

The policy implications of the analysis proposed could be relevant for the evolution of the WTO DOHA Development Round. A major focus on services, in fact, could help expanding the "policy space" faced by the negotiators, possibly increasing the likelihood of a successful conclusion of the round. Moreover, this paper inform also the recent debate about the need of a revaluation of the yuan. Allowing the U.S. to increase its exports of services (not necessarily to China) might help alleviating global imbalances even without movements in the exchange rates.

The second essay is titled **"Estimating Trade and Investment Flows: Partners and Volumes"**. I present empirical evidence from a large sample of countries for the period 2000-2006. Bilateral foreign direct investment (FDI) flows are almost never observed in the absence of bilateral trade flows, thus configuring an order of trade and investment flows. I document a similar pattern using bilateral foreign affiliate sales (FAS), aggregating them up from a large firm level dataset (ORBIS), which includes over 45,000 firms.

I propose a model where heterogeneous firms face a proximity-concentration tradeoff when they decide whether to serve foreign markets through export or FDI. I derive theory-based gravity-type equations for the aggregate bilateral trade and foreign affiliate sales (FAS) flows. I then suggest a two-stage estimation procedure.. In the first stage, a ordered Probit model is used to retrieve consistent estimates of the terms needed to correct the flows equations for heterogeneity and selection. In the second stage, a maximum likelihood estimator is applied to the corrected trade and FAS equations.

The main results of the analysis are as follows: 1) The impact of distance, border and regional trade agreements on bilateral foreign affiliate sales becomes substantially smaller after controlling for selection and firms' heterogeneity (hence separating the impact on the extensive versus the intensive margin). 2) The same "attenuation" result is found also for the trade equations, consistently with HMR. 3) When FAS are observed, failing to take this into account when correcting for heterogeneity and selection in the trade equations leads to differences in the estimated coefficients.

The third essay is titled **"Some Evidence on the Importance of Sticky Wages"**, and is co-authored with Susanto Basu and Peter Gottshalk.

Nominal wage stickiness is an important component of recent medium-scale structural macroeconomic models, but to date there has been little microeconomic evidence supporting the assumption of sluggish nominal wage adjustment. We present evidence on the frequency of nominal wage adjustment using data from the Survey of Income and Program Participation (SIPP) for the period 1996-1999. The SIPP provides high-frequency information on wages, employment and demographic characteristics for a large and representative sample of the US population.

The main results of the analysis are as follows. 1) After correcting for measurement error, wages appear to be very sticky. In the average quarter, the probability that an individual will experience a nominal wage change is between 5 and 18 percent, depending on the samples and assumptions used. 2) The frequency of wage adjustment does not display significant seasonal patterns. 3) There is little heterogeneity in the frequency of wage adjustment across industries and occupations 4) The hazard of a nominal wage change first increases and then decreases, with a peak at 12 months. 5) The probability of a wage change is positively correlated with the unemployment rate and with the consumer price inflation rate.

To a certain extent, the three essays presented here are self-contained and deal with three different issues regarding international economics and macroeconomics. Going to a deeper level, however, the essays are linked by a common feature: they are three examples of economic research across fields. The first essay, in fact, is an example of the growing fields at the edge between international trade and international macroeconomics. While the trade of goods and services and the dynamics of macroeconomic variables such as the current account are highly interconnected in the real world, these two fields have been characterized by a large divide in the last thirty years in the economic literature. The second essay is an example of a joint study of international trade and investment flows. Also in this case, while conceptually clearly interconnected, these topics have been usually studied separately by the economic literature. Finally, the third essay is an example of research across fields (labor economics and macroeconomics) and techniques (micro-level analysis informing macroeconomic models). In this last case, macroeconomists were interested in estimating certain wage dynamics parameters highly used in macro models. However, they were largely unaware of the fact that labor economists had the data to answer those research questions. On the other hand, the labor economists had the data, but not the questions. I hope that these essays might help increasing further the awareness that more communication between economists working in different fields can bring to valuable insights.

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Contents

Chapter 1

Comparative Advantage, Service Trade, and Global Imbalances

1.1 Introduction

The accumulation of current account deficits in the U.S., accompanied by the corresponding surpluses registered in Japan, Germany, China and other countries has generated the phenomenon known as "global imbalances," described as "probably the most complex macroeconomic issue facing economists and policy makers" (Blanchard and Milesi-Ferretti, 2009). The emergence of such imbalances has also been recently suggested as one of the sources of the financial and economic crisis that began in 2007 (Obstfeld and Rogoff, 2009; Bernanke, 2009).

The motivation for this paper is best understood by exploring the composition of global imbalances. Figure 1.1 shows the U.S. current account disaggregated into its two main components: the trade balance and the income balance.[1] Clearly, the current account deficit tracks the trade balance very closely, indicating that trade might play an important

[1]The current account is net of transfers. The trade balance is the difference between the value of exports of goods and services and the imports of goods and services. The income balance is the difference between the payments from abroad and the payment to the foreigners for i) profits from fdi, ii) returns from portfolio investments (equities and bonds), and iii) interests on government bonds.

role in global imbalances.

Further disaggregating the U.S. current account deficit reveals an interesting fact. Figure 1.2 shows the further disaggregation of the trade balance into its two components: the goods balance and the service balance. It is clear from the picture that the U.S. trade imbalance is due to a large deficit in the goods balance and a modest surplus in the service balance.

Figure 1.3 shows a similar disaggregation of Japan's current account surplus. While there are visible fluctuations, Japan's large trade surplus is the combination of a surplus in the goods balance and a deficit in the service balance. The same is true for Germany (Figure 1.4) and China (Figure 1.5).

Starting from this stylized fact, I propose a new explanation for the formation of global imbalances. The focus of this story is the interplay between the comparative advantage of the U.S. in services and the asymmetric trade liberalization process in goods trade versus service trade that took place in the last two decades, particularly since the mid-nineties.

The conclusion of the Uruguay Round and the advent of the World Trade Organization (WTO) in 1995 spurred the liberalization of trade, especially in goods and agricultural products. At the same time, the General Agreement for Trade in Services (GATS) was signed. However, after more than fifteen years, the liberalization process in service trade does not seem to have made much progress (Adlung, 2009).

The first part of the paper establishes empirically the presence of a revealed comparative advantage of the U.S. in services and the emergence of a negative relation between specialization in export of services and current account for a large sample of countries, starting in the middle of the nineties. A simple Balassa-type index of revealed comparative advantage (RCA) confirms that the U.S. is relatively specialized in services, while Germany, Japan and China display a revealed comparative advantage in goods. Moreover, I document a systematic negative relation between revealed comparative advantage in services and current account for a large sample of OECD and developing countries. This relation is strongly statistically significant (both in a cross-section and in a panel regression analysis) and robust after controlling for standard determinants of the current account as well as controlling for financial development. Past revealed comparative advantage is able to explain about 50% of the variation of current account balances. Importantly, this relation emerges only when considering the post-1995 period.

In the second part of the paper, I use the structural gravity model developed by Anderson and van Wincoop (2003) to document the presence of an asymmetry in the liberalization of goods trade versus service trade. I use the newly proposed concept of the constructed home bias index (CHB) developed by Anderson and Yotov (2010a) to quantify the extent of this asymmetry. The CHB index is the ratio of the realized internal trade in a given sector relative to the internal trade that would prevail in a frictionless world. This index is a pure number, so it can be compared across different sectors and an appropriately weighted average of this index captures the liberalization process in manufacturing versus service trade at the world level. While the index for manufacturing trade, available from 1994, is declining since the mid-nineties, the index for services is virtually flat. The extent of the asymmetry is of the order of 15% over the 12 years for which data are available.

In the third part of the paper, I show that a simple two-period model can rationalize the emergence of current account deficits in the presence of such an asymmetric liberalization process. The structure of the model is minimal - endowment economies with complete specialization - to maximize transparency of the mechanisms and results. The anticipation of a future reduction of impediments to trade in services generates an increase in savings in the service-importing countries (like Japan, Germany, and China) due to an increase in the consumption-based real interest rate (intuitively, it is convenient to save for future times when services will be available at cheaper prices). At the same time, the anticipation of future higher income and a relatively more expensive consumption basket generates a deficit in the service-exporting countries (like the U.S. or the UK) due to the combination of a lower consumption-based real interest rate and a positive wealth effect (intuitively, it is convenient to enjoy increased consumption of cheap manufactured goods today by borrowing against a future increase in income from higher future demand for services). These results hold both in a small-economy (partial equilibrium) version of the model and in a two-country (general equilibrium) version, provided that the intertemporal elasticity of substitution is sufficiently large. I solve the log-linear version of the two-country model analytically, and I show how the current account responds to changes in trade costs when these changes are asymmetric, thus showing that trade policy can indeed affect current account balances in dynamic settings where expectations play an important role.

Finally, I evaluate the quantitative relevance of this explanation for global imbalances. I develop a multi-period version of the model and solve it (in levels) under perfect foresight, using as the exogenous driving forces a profile of reduction of trade costs analogous to what found in the empirical analysis. If the model is shocked with a symmetric trade liberalization process in both manufacturing and service trade, there is no change in the net foreign asset position of either country. When the model is shocked with the actual asymmetric liberalization process found in the data, it generates a current account deficit in the service-exporting country of about 1% of GDP, roughly 20% of U.S. current account at the peak of global imbalances (2006).

This paper is linked to several strands of the literature. First, it is related to the many explanations for global imbalances that have been proposed. After Bernanke (2005) proposed the "savings glut" hypothesis, influential papers have emphasized the role of heterogeneity across countries in the ability to supply assets to savers (Caballero, Farhi and Gourinchas, 2008) or in different levels of financial development (Mendoza, Quadrini and Rios-Rull, 2009) as the key elements to global imbalances. I refer to these theories as "financial" explanations. Kamin and Gruber (2009), however, find little evidence that differences in financial sector development can explain the cross-sectional pattern of external balances. The explanation proposed in this paper is "real," in that it focuses on the consequences of comparative advantage and (asymmetric) trade liberalization across goods and services.

This is not the first "real" story for global imbalances. Ju and Wei (2009) and Jin (2009) also explore trade channels that can potentially explain the emergence of external imbalances. Comparative advantage plays an important role, but the economic channels in these papers operate via the production side, while the key mechanisms of the model presented in this paper operate through the consumption side.

Other "real" explanations for global imbalances include the anticipation of a rising U.S. future share in world output (Engel and Rogers, 2006) and the productivity dynamics in the non-traded sector (Cova et all, 2005). Other stories for global imbalances include the "dark matter" argument (Hausmann and Sturzenegger, 2005) and the Bretton Wood II hypothesis (Dooley, Folkerts-Landau, and Garber, 2003).[2] Blanchard and Giavazzi (2002),

[2]As pointed out by the authors, the Bretton Wood II Hypothesis relates more to the reason why Asian countries are willing to keep financing U.S. current account deficits than to the reason that determines the deficits themselves.

though not directly addressing the issue of global imbalances, find that greater real and financial integration can exacerbate the deficits of borrowing countries and increase the surpluses of creditor countries in the presence of growth differentials. Finally, there are also "behavioral" explanations for global imbalances, which include the asset prices boom of the late nineties and the associated perception of greater wealth by the American consumers (Laibson and Mollerstrom, 2010). None of these explanations focus on the service sector and the asymmetric trade liberalization of goods trade versus service trade.

Second, the paper is linked to a vast literature on service trade. Mattoo, Stern, and Zanini (2008) and Francois and Hoekman (2010) provide an excellent overview. Mann (2004) provides an empirical analysis of the implications of service trade for the U.S. current account balance.

Finally, this paper is related to the literature that studies international macroeconomic models in the presence of trade costs. After the seminal contribution by Obstfeld and Rogoff (2001), several have explored the implications of introducing trade costs into international macroeconomic models. Coeurdiacer (2009) is an example from which I borrow some modeling choices for the two-period model.

The paper is structured as follows. Section 1.2 contains further motivating evidence. Section 1.3 presents the empirical analysis. Section 1.4 presents the two-period model in its partial equilibrium version. Section 1.5 presents the two-country model. Section 1.6 presents the multi-period model and the quantitative exercise. Section 1.7 concludes.

1.2 Motivating Evidence

The first ingredient of the story for the emergence of global imbalances proposed in this paper is the U.S. comparative advantage in services. The aim of this section is to provide empirical evidence of it as well as to document a new stylized fact, namely the emergence from the mid nineties of a strong negative relation between revealed comparative advantage in services and the current account.

Before proceeding, a disclaimer is appropriate. Obviously, the service sector is extremely heterogenous. It encompasses very different activities such as construction services, transport services, financial services, insurance services, business services (such as legal, architectural, R&D, and advertisement services), educational services, and health services.[3] I keep an aggregate perspective here partly for consistency with the models that will be introduced in the following sections and partly because of data limitations.[4]

1.2.1 The U.S. Comparative Advantage in Services

I document the comparative advantage of the U.S. in the production and export of services using data from the World Development Indicators of the World Bank, which include extensive coverage of export and import of goods and services for a large sample of countries of which I pick 83.[5] The simplest way to assess the presence of comparative advantage is the revealed comparative advantage index (RCA), first introduced by Balassa (1965). The RCA index expresses the ratio of the export share of a given sector in country

[3]See Marchetti and Roy (2008) for a case-study approach.

[4]Repeating the exercises of the next subsections at a more disaggregated level would certainly represent a fruitful avenue for future research, even though the data constraints would likely become more stringent.

[5]I'm using in this section a sample of 83 countries: the OECD countries and the upper-middle income countries according to the World Bank classification, plus China, India, Indonesia, Singapore and Saudi Arabia

i relative to the export share of that sector in the world as a whole. Thus, the RCA is an index of relative export specialization and can be expressed as follows:

$$RCA_{i,k} = \frac{\frac{EXP_{i,k}}{\sum_k EXP_{i,k}}}{\frac{EXP_{WLD,k}}{\sum_k EXP_{WLD,k}}},$$

where i is a country and $k = \{Goods, Services\}$. A value of the RCA larger than 1 for a given sector k indicates that the country is relatively specialized in export in that sector, thus revealing the presence of a comparative advantage of the country in that sector.

Figure 1.6 reports the evolution of the RCA index in services for the U.S. and three other countries (Japan, China, and Germany), which are among the largest creditors of the U.S. (they represent on average 52% of the total U.S. current account deficit for the period 2000-2008 and 63% in 2008). The RCA in services for the U.S. is larger than 1 and rising over time. On the other hand, the RCA indexes for services for China, Germany, and Japan are all smaller than 1. In the case of China, the RCA displays a clear downward trend. Germany and Japan do not display clear trends.

Not surprisingly, since I am considering a world where only two "goods" are produced and traded (goods and services), the picture for the RCA in goods (not shown) looks symmetrical, with the U.S. displaying an increasing revealed comparative disadvantage in the export of goods. On the other hand, Japan, Germany, and China report RCA indexes above 1 (and rising, in the case of China).

An important feature of Figure 6 is the relative stability of the indexes. In fact, out of the 83 countries of the sample, only in very few cases the index cross the threshold of one in the period considered. Interestingly, two exceptions are India and Ireland, which became increasingly specialized in the export of services only in recent times. This feature is relevant for the models that will follow, where comparative advantage will be assumed as exogenously given and not time varying.

1.2.2 A New Stylized Fact

Figure 1.7 and 1.8 report scatter plots relating the current account over GDP to the revealed comparative advantage in services a decade before. The sample include the OECD countries plus the BRICS (Brazil, Russia, India, China and South Africa). Figure 1.7 refers to 1995 while figure 1.8 to 2006.[6] As it is clear from the pictures, a strong negative relation between comparative advantage in services and current account is found to be strong in 2006, but virtually absent in 1995. Figure 1.9 reports the same scatter plot as the one in figure 1.8, but extends the sample to include also the upper-middle income countries. The negative relation between specialization in services and current account is striking.

In order to validate this visual evidence in a regression framework, table 1.1 reports the result of a cross-section regression using as dependent variable the current account over GDP ratio. In the first three columns, the reference year is 2006 while in the second three columns the reference year is 1995. Comparing the first and the third column, one can see how the strength of the negative relation between past specialization in services and the current account is much stronger in 2006 (the coefficient is seven times bigger) and the explanatory power is much higher (the R^2 is 50% in 2006 versus 8% in 1995). Columns two, three, five and six just confirm that this negative relation is robust after controlling for real gdp per capita, gdp growth, openness and (importantly) a measure of financial

[6]The reason for choosing 1995 is the entrance into force of the WTO, 2006 was instead the year in which global imbalances peaked, before the advent of the Great Recession.

development.

Given the usual limitation of cross-sectional analysis, in table 1.2 I present a panel analysis. The data are organized in 5-years non overlapping averages and span the period 1970-2010. I test the presence of a correlation between the average current account over GDP ratio and the average specialization in service trade in the previous 5-years period. All the regressions include time and country fixed effects. As table 2 shows clearly, a strong negative relation is found between past specialization in services and the current account, but only for the post 1995 period. This relation is robust after controlling for the other determinants of current account used in the cross-sectional analysis. Interestingly, the proxy used for financial development (domestic credit over GDP) displays a negative correlation with current account for the pre-1995 period, but the relation becomes statistically insignificant in the post-1995 period.[7].

The evidence and the theoretical mechanisms proposed in the following sections to explain the emergence of global imbalances are broadly consistent with this new intriguing stylized fact.

1.3 Trade Liberalization in Service Trade and Manufacturing Trade

The second main ingredient of the story for global imbalances proposed in this paper is the existence of an asymmetric trade liberalization process in goods trade versus service trade starting from the mid-nineties. The aim of this section is to provide empirical evidence

[7]A result that is reminiscent of Gruber and Kamin(2009).

of such asymmetry.

Establishing qualitatively a gap in the liberalization of trade in services compared to manufacturing trade is not difficult. If one wants to take a long-term perspective regarding international negotiations, for instance, one would immediately realize that the negotiations on liberalization of trade in goods started in the late forties with the General Agreement on Trade and Tariffs (GATT). Only in 1995, however, a negotiating framework for the liberalization of service trade was established with the GATS (General Agreement for Trade in Services), together with the institution of the World Trade Organization (WTO).

Even restricting the attention to the last fifteen years, however, reveals that little progress has been made in the liberalization of service trade. A recent WTO study, for example, concludes that "there has been virtually no liberalization under the GATS to date" (Adlung, 2009). Moreover, as stressed by Mattoo and Stern (2008), "in the negotiation under the Doha Development Agenda, services have received surprisingly little attention."

Trying to quantify the extent of this asymmetry is a much more difficult task. As stressed by Deardorff and Stern (2008), while barriers to trade can be summarized in "tariff equivalent" measures in the case of trade in goods, things are more complicated for the case of services, since such simple measures do not exist. Barriers to service trade usually take the form of regulations, which are inherently more difficult to measure.

Moreover, services can be provided abroad through different modes. The GATS classifies four typical modes of provisions of services:

1. *Cross Border (mode 1).* Examples could be a software service provided through e-mail

or a call center placed abroad.

2. *Consumption Abroad (mode 2).* Examples of this mode of provision are foreign students consuming education services or tourism.

3. *Commercial Presence (mode 3).* Examples could be a financial firm or a consulting firm opening an affiliate in a foreign country through a foreign direct investment (FDI).

4. *Presence of Natural Person (mode 4).* For example, a doctor moving to perform surgery in a different country or the employees of a multinational enterprise (MNE) moving on a temporary basis.

Not surprisingly, different types of possible restrictions correspond to different types of provision mode.

As Deardorff and Stern (2008) report, researchers have used a variety of methods to estimate empirically the barriers to trade in services. The most popular methods have been the use of indexes of restrictiveness, price-impact or quantity-impact studies in reduced form regressions, and the gravity model studies.[8] In what follows, I propose a novel approach based on the structural gravity model and the constructed home bias (CHB) index developed by Anderson and Yotov (2010a).

The main advantage of the approach will be to give quantitative measures of trade restrictions that are conceptually comparable for manufacturing trade and service trade and are also time varying. The main disadvantage will be that, as all the indirect measures of trade restriction, these may include the effects of other (unidentified) frictions.

[8]See Deardorff and Stern (2008) and references therein.

1.3.1 Structural Gravity

The structural gravity model is extremely successful at fitting bilateral trade data. Following Anderson and van Wincoop (2003), let X_{ij}^k be the total shipment from the origin country i to the destination country j in sector k, E_j^k the total expenditure in sector k in the destination country j, and Y_i^k the total output of sector k in the origin country i. The structural gravity model can be expressed as follows:

$$X_{ij}^k = \frac{Y_i^k E_j^k}{Y^k} \left(\frac{t_{ij}^k}{P_j^k \Pi_i^k} \right)^{1-\theta_k}, \tag{1.1}$$

$$\left(\Pi_i^k\right)^{1-\theta_k} = \sum_j \left(\frac{t_{ij}^k}{P_j^k} \right)^{1-\theta_k} \frac{E_j^k}{Y^k}, \tag{1.2}$$

$$\left(P_j^k\right)^{1-\theta_k} = \sum_i \left(\frac{t_{ij}^k}{\Pi_i^k} \right)^{1-\theta_k} \frac{Y_i^k}{Y^k}. \tag{1.3}$$

in (1) Y^k represents the world output of sector k and t_{ij}^k represents the bilateral trade cost of shipping a unit of sector k good from country i to country j. P_j^k and Π_i^k are the inward and outward multilateral resistance terms. P_j^k summarizes the average resistance to inward trade of country j, as if country j were importing all the goods from a theoretical "international" market. Similarly, Π_i^k represents the average outward resistance to trade of country i, as if country i were exporting all the goods to the same theoretical "international" market.[9] θ_k is the elasticity of substitution between goods in a given sector k.

Interestingly, the structural gravity model contains a prediction for the amount of bi-

[9] Intuitively, what really matters are the bilateral trade costs relative to the average difficulty to trade. As an example assume that trade costs depend only on distance and take the example of Australia and New Zealand. While Australia and New Zealand are far from each other in absolute term, both of them are very far from all their other trading partners. This makes them "relatively close".

lateral trade that would prevail in a frictionless world. If $t_{ij}^k = 1$ for every country pair ij, in fact, $\Pi_i^k = P_j^k = 1$, and $X_{ij}^k = \frac{Y_i^k E_j^k}{Y^k}$.

The data source for the analysis is the Service Trade Database, developed by Francois et al (2009), which contains data on bilateral service trade flows for a large sample of countries for the period 1999-2005. Manufacturing trade data are from UN-COMTRADE. A data appendix contains more details.[10]

I estimate the gravity equation (1.1) for aggregate manufacturing and service trade in a sample of 23 OECD countries plus China[11] for the period 1994-2005. For Services I can only study the period 1999-2005 because of data availability.

I estimate equation (1.1) in levels by generating a transformed dependent variable $dep_k = \frac{X_{ij}^k Y^k}{Y_i^k E_j^k}$ and then using Poisson pseudo maximum likelihood (PPML) as suggested by Santos-Silva and Tenreyro (2006).[12] I use standard proxies for bilateral trade costs: population-weighted distance, dummies for common border, common language, common colonial past, common legal origin, and a dummy that is equal to 1 in case of internal trade.[13] I include also exporter and importer fixed effects. Table 1.3 reports the results obtained for different years. The first two columns refer to manufacturing, the second two columns to total services and the third two columns to the a sub-sample of strictly tradeable services. I include in this sub-sample only business services (for instance consulting, legal or accounting services etc.), financial services and transportation services.[14]

[10]From now on I will focus on manufacturing trade as opposed to goods trade and exclude agricultural products and commodities. The reason is that in order to compute the internal trade I will need to include also gross output data, which are more readily available for manufacturing than for primary sectors.

[11]The reduction in the number of available countries is imposed by the need of gross output data. See the data appendix for details.

[12]I also used the simple OLS regressions for comparison. While the points estimates are inevitably different, the same qualitative results hold.

[13]X_{ij}^k is defined as gross output value minus exports

[14]In the U.S. the category of tradeable services accounts for about 30% of total service output. When

Three observations stand out. First, physical distance appears to be a more important determinant of manufacturing trade than service trade (and in both cases, its importance rises slightly over time). This is not surprising, considering the evolution of technology in the delivery of services across border. Second, the presence of a common colonial origin and a common legal system appear to be more important for service trade than manufacturing trade. The coefficients tend to fall over time, but more so in the manufacturing regressions than in the service ones. Third, the coefficient on the dummy for internal trade is larger for service trade than manufacturing trade. Moreover, consistent with intuition, the coefficient is slightly lower when considering only the sub-sample of strictly tradeable sectors. In both sectors, the coefficients decline over time, but also in this case more so in the manufacturing sector than in the service sector.

1.3.2 The CHB Index: Asymmetric Trade Liberalization

Taking the estimated $\left(t_{ij}^{k}\right)^{1-\theta_k}$ of the first stage regressions, I can solve for the multilateral resistance terms P_j^k and Π_i^k using equations (2) and (3). To do so, I need to impose the normalization that $P_i^k = 1$ for each k, where i is a convenient country (in this case, Austria). After doing that, I finally can compute a constructed home bias index, equal to the ratio of internal trade to the internal trade that would prevail in a frictionless world:

$$CHB_{ik} = \left(\frac{t_{ii}^{k}}{P_i \Pi_i}\right)^{1-\theta_k}. \tag{1.4}$$

This index has the important property of being invariant to the elasticity of substitution

using this sub-sample of service industries, I cannot include Poland, New Zealand and China in the analysis due to lack of output data.

θ_k.[15] This is an important reason to use the CHB to proxy for barriers in service trade. While estimates of the elasticity of substitution are available for several manufacturing sectors after the work by Broda, Greenfield and Weinstein (2006), to the best of my knowledge there are no reliable estimates for the service sector.[16] Table 1.4 reports the results for the CHB in manufacturing and the service sector at the level of the single country for selected years.

The U.S. appears to be the most "open" country in the sample, in both manufacturing and services. Interestingly, the U.S. is the only country where the CHB in services is smaller than the CHB in manufacturing. In all the other countries the CHB in the service sector is higher than in manufacturing.

As for the time dimension, in the U.S. the CHB index is roughly constant both in the manufacturing sector and in the service sector. Other countries present very different patterns. In Germany, for example, the CHB grows slightly over time in the service sector and has an inverted-u shape in manufacturing. Japan is characterized by increasing levels of CHB in both manufacturing and services. China presents a significant reduction of the CHB in manufacturing (more than 60% between 1994 and 2005) and services (around 50% from 1999 to 2005). The levels of the Chinese CHB indexes, however, are several times higher the corresponding figures for economies the U.S., Germany and Japan.

In order to have a global CHB index for manufacturing and services, I aggregate the results obtained for individual countries. Following Anderson and Yotov (2010b), I use as weights the frictionless internal trade shares. I'm hence able to express a world-level CHB

[15]The reason is that the first stage regression identified $\left(t_{ij}^k\right)^{1-\theta_k}$, while solving the system (2)-(3) literally solves for $\left(\Pi_i^k\right)^{\theta_k-1}$ and $\left(P_i^k\right)^{\theta_k-1}$.

[16]This clearly represents another interesting avenue for future research. Of course, disclaimer about the level of heterogeneity in the service sector applies here as well.

index:

$$CHB_k = \sum_i \left(\frac{t_{ii}^k}{P_i \Pi_i} \right)^{1-\theta_k} \frac{\frac{Y_i^k E_i^k}{Y^k}}{\sum_i \frac{Y_i^k E_i^k}{Y^k}}. \tag{1.5}$$

Figure 1.10 reports the evolution of the CHB in manufacturing and services (considering both total services and the sub-sample of tradeable services). Two features stand out. First, the average level of CHB is higher in services than in manufacturing. Intuitively, the index obtained considering total services is higher than the one obtained using only tradeable services. Second, the CHB in services is almost entirely flat over the 1999-2005 period. On the contrary, the CHB in manufacturing is declining over the period 1994-2005, with an initial acceleration around 1996 (the year after the advent of the WTO).

Figure 1.11 normalizes the CHB indexes to 100 at the beginning of the period.[17] From Figure 11, it is possible to quantify the asymmetry of the two liberalization processes to be roughly 15% of the initial values over the period considered.[18]

1.3.3 Caveats

Before proceeding, it is useful to qualify the results above with some caveats. The first concerns data availability. The Trade in Services Database (TSD) contains information from balance of payments statistics. This means that the trade data considered cover only modes 1 and 2 of provision of services. While mode 4 is almost irrelevant quantitatively, mode 3 (FDI) is certainly a very important way of providing services to another country. There is some coverage of this mode of provision in the TSD, but not the bilateral data

[17] 1994 for manufacturing and 1999 for services.

[18] The implicit assumption is obviously that the pattern of the CHB for services is flat in the period 1994-1999. While this appears a reasonable assumption, It is important to stress that it is an assumption.

necessary to perform the analysis of the previous sections. Francois et all (2009) suggest that mode 3 accounts for roughly 35% of global service trade. In the case of the U.S., this share is larger (around 60%). So it is important to know that the analysis of the previous section miss an important part of the action.[19] However, data on bilateral foreign affiliate sales in the service sector are virtually non-existent.[20] Presumably, taking into account the sales of foreign affiliates would indicate a faster pace of liberalization in services than found in the previous section.

Second, the sample is limited to 23 OECD countries plus China. The reason for this limitation is the need of gross output data for manufacturing and services, which are not available for most countries. Exanding the sample of countries to include more developing countries would likely generate a stronger asymmetry in the pace of liberalization of manufacturing versus service trade. These extensions of the empirical analysis will be feasible when more data are available.

[19] In 2008, the U.S. total export of services was 518.3 billions while the total sales of U.S. foreign affiliates in the service sector were 761.5 billions

[20] they exists only at the aggregate level. Only the US records the data also by partner country, but bilateral data are needed in order to use a structural gravity model

1.4 A Simple Two Period Model: Partial Equilibrium

In this section, I develop a simple theoretical model that explains the channels though which the interaction between comparative advantage and asymmetric trade liberalization can explain the emergence of current account deficits. I start with a partial equilibrium setting in this section and move to general equilibrium in the next section.

A small open economy is populated by a representative household that lives for two periods with perfect foresight. Two goods are consumed: a home good (h) and a foreign good (f). The endowment of the home good for the two periods is Y_t^h with $t = \{1, 2\}$. The prices of the home goods is p_t^h. The foreign good is imported from the rest of the world at exogenous price p_t^{f*}, set to be the numeraire. The internal price of the imported good is $p_t^f = \tau_t^f p_t^{f*}$, where $\tau_t^f > 1$ is an iceberg trade cost that captures impediments to trade.

The household maximizes:

$$\frac{C_1^{1-\frac{1}{\sigma}} - 1}{1-\frac{1}{\sigma}} + \beta \frac{C_2^{1-\frac{1}{\sigma}} - 1}{1-\frac{1}{\sigma}},$$

where $\sigma > 0$ is the intertemporal elasticity of substitution and $\beta \in (0, 1)$ is the discount factor.

The consumption basket aggregates the home goods and foreign goods with unitary elasticity of substitution:

$$C_t = \left(\frac{C_t^h}{\nu}\right)^{\nu} \left(\frac{C_t^f}{1-\nu}\right)^{1-\nu}, \qquad 0 \leq \nu \leq 1$$

The assumption of unitary elasticity of substitution allows me to obtain transparent closed-form solutions. The price index associated to the consumption basket can be written:

$$P_t = \left(p_t^h\right)^{\nu} \left(\tau_t^f p_t^{f*}\right)^{1-\nu}.$$

The representative household has access to a riskless saving instrument denominated in units of the world currency that yields a constant gross return $(1+r^*)$. The budget constraints for period 1 and 2 imply the intertemporal budget constraint:

$$P_1 C_1 + \frac{P_2 C_2}{1+r^*} = p_1^h Y_1^h + \frac{p_2^h Y_2^h}{1+r^*}. \tag{1.6}$$

The household's maximization problem allows to express the following Euler equation for consumption:

$$C_1 = \beta^{-\sigma} \left((1+r^*) \left(\frac{P_1}{P_2} \right) \right)^{-\sigma} C_2 \tag{1.7}$$

Equations (1.6) and (1.7) make it possible to solve for C_1 and C_2. In particular:

$$P_1 C_1 = \frac{1}{1 + \beta^{\sigma} (1+r^*)^{\sigma-1} \left(\frac{P_1}{P_2} \right)^{\sigma-1}} \left(p_1^h Y_1^h + \frac{p_2^h Y_2^h}{1+r^*} \right). \tag{1.8}$$

The value of consumption in period 1 is a share of lifetime income. The share depends on the discount factor β, the intertemporal elasticity of substitution σ and the consumption-based real interest rate $(1+r^*) \frac{P_1}{P_2}$. Since the model does not feature investment, the current account in period 1 equals savings:

$$CA_1 = p_1^h Y_1^h - \frac{1}{1 + \beta^{\sigma} (1+r^*)^{\sigma-1} \left(\frac{P_1}{P_2} \right)^{(\sigma-1)}} \left(p_1^h Y_1^h + \frac{p_2^h Y_2^h}{1+r^*} \right). \tag{1.9}$$

Equation (9) is the starting point for understanding the possible effects of asymmetric liberalization in trade of the two goods.

1.4.1 Asymmetric Trade Liberalization

Suppose that the two goods are "manufactured goods" (m) and "services" (s). A simple way to capture an asymmetric trade liberalization process is represented in Figure 1.12. At the beginning of period 1, impediments to trade in manufactured goods fall and stay low for both periods (dashed line). At the same time, it is known that the impediments to trade in services will fall only at the beginning of period 2 (solid line).

1.4.2 The Service-Importing Countries Perspective

Take now the perspective of service-importing countries (like Germany, Japan, or China) and consider the effect of the experiment described above. Adjusting appropriately[21], the ratio of the price indexes between period 1 and 2 becomes:

$$\frac{P_1}{P_2} = \left(\frac{p_1^m}{p_2^m}\right)^{\nu} \left(\frac{\tau_1^s}{\tau_2^s}\right)^{1-\nu} = \left(\frac{\tau_1^s}{\tau_2^s}\right)^{1-\nu}, \tag{1.10}$$

where the second equality holds under the assumption that p_t^m is constant. Substituting equation (10) into equation (9) yields the following expression for the current account in period 1:

$$CA_1 = p_1^m Y_1^m - \frac{1}{1 + \beta^{\sigma}\left[(1+r^*)\left(\frac{\tau_1^s}{\tau_2^s}\right)^{1-\nu}\right]^{\sigma-1}} \left(p_1^m Y_1^m + \frac{p_2^m Y_2^m}{1+r^*}\right) \tag{1.11}$$

Inspection of equation (1.11) shows that $\frac{\partial CA_1}{\partial \tau_2^s} < 0$ if $\sigma > 1$. Anticipating a decrease in trade costs for services in the future pushes households to consume less and save more today (i.e., to run a current account surplus). The channel is best explained by Figure 1.13, where

[21] The home good here is manufacturing, m, and the foreign good is services, s.

Z1 and Z2 are the endowments in the two periods in units of the consumption basket.[22] Suppose the starting point is a situation where preferences and endowment are such that the small open economy would not save nor borrow (so where consumption for periods 1 and 2 is at point Z, coinciding each period with the endowment). A decrease in τ_2^s represents an increase in the consumption-based real interest rate. If the intertemporal elasticity of substitution is sufficiently large, this induces the households to delay consumption to enjoy lower prices in the future (so the consumption point becomes C while the endowment point is still Z).

Notice that the assumption $\sigma > 1$ is crucial for obtaining this result in this partial equilibrium context. While usually the empirically relevant values of σ are thought to be smaller than 1, recent research based on micro-level data found values of σ as high as 2 (Gruber, 2005).[23] I relax the assumption $\sigma > 1$ in the two-country general equilibrium model.

1.4.3 The Service-Exporting Countries

Take now the perspective of the service-exporting countries (like the U.S. or UK). The ratio of the price indexes in the two periods can now be expressed as:[24]

$$\frac{P_1}{P_2} = \left(\frac{p_1^s}{p_2^s}\right)^{\nu} \left(\frac{\tau_1^m}{\tau_2^m}\right)^{1-\nu} = \left(\frac{p_1^s}{p_2^s}\right)^{\nu}, \tag{1.12}$$

where the second equality follows from the fact that the trade cost for manufacturing is constant across periods. For these countries, the analogue to equations (1.9) and (1.11) is:

[22] Formally, $Z1 = \frac{p_1^m}{P_1} Y_1^m$ and $Z2 = \frac{p_2^m}{P_2} Y_2^m$

[23] Moreover, Barro (2009) explains how $\sigma > 1$ is actually needed to explain features of asset price dynamics.

[24] Remember that now the Home good are services and the foreign good is manufacturing

$$CA_1 = p_1^s Y_1^s - \frac{1}{1+\beta^\sigma \left[(1+r^*)\left(\frac{p_1^s}{p_2^s}\right)^\nu\right]^{\sigma-1}} \left(p_1^s Y_1^s + \frac{p_2^s Y_2^s}{1+r^*}\right). \tag{1.13}$$

A reduction in the trade costs of exporting services in the future can generate an increase in the future internal price of services if the increase in demand abroad is strong enough. With $\sigma > 1$, an increase in p_2^s unambiguously generates a current account deficit at period 1. Figure 1.14 explain the economic channels, here a combination of a decrease in the consumption-based real interest rate and a positive wealth effect coming from the increase in the value of the endowments.[25] Starting from a point like Z, where no international borrowing occurs, the decrease in the consumption-based real interest rate is making the budget constraint flatter while the wealth effect is increasing the value of future endowments. As a result, the new optimal allocation for consumption is C, and the country runs a current account deficit.

A natural question is whether the simple channels illustrated in Figures 1.13 and 1.14 would survive a more comprehensive treatment that considers the general equilibrium effects of the change in trade costs on all prices. To address this question, I illustrate the results of a comparably simple two-country, general equilibrium model of the world economy.

1.5 General Equilibrium

The world consists of two countries : Home and Foreign (with foreign variables denoted by *). Each country is populated by a representative household that lives for two periods. Two goods are consumed: a home good (services, s) and a foreign good (manufactured

[25] Here $Z1 = \frac{p_1^s}{P_1} Y_1^s$ and $Z2 = \frac{p_2^s}{P_2} Y_2^s$

goods, m). The endowment of the home good is Y_t^s with $t = \{1, 2\}$. The endowment of the foreign good is Y_t^{m*} with $t = \{1, 2\}$. The price home good at Home is p_t^s. The price of the home good in Foreign is $p_t^{s*} = \tau_t^s p_t^s$, where $\tau_t^s > 1$ is an iceberg trade cost. The foreign good m is imported in Home from Foreign. The Home price of the foreign good is $p_t^m = \tau_t^m p_t^{m*}$, where p_t^{m*} is the price of the foreign good in Foreign (set to be the numeraire) and $\tau_t^m > 1$ is an iceberg trade cost.

In both countries, households maximize lifetime utility, given by:

$$\frac{X_1^{1-\frac{1}{\sigma}}-1}{1-\frac{1}{\sigma}} + \beta \frac{X_2^{1-\frac{1}{\sigma}}-1}{1-\frac{1}{\sigma}}$$

where $X = C$ or C^* depending on the country. The asset menu features only an international bond denominated in units of a common world currency. The first-period and second-period budget constraints are, respectively:

$$B_1 = p_1^s Y_1^s - P_1 C_1, \qquad B_1^* = Y_1^{m*} - P_1^* C_1^*, \tag{1.14}$$

$$P_2 C_2 = p_2^s Y_2^s + (1 + r_1) B_1, \qquad P_2^* C_2^* = Y_2^{m*} + (1 + r_1) B_1^*, \tag{1.15}$$

where B_1 and B_1^* are the net bond positions of Home and Foreign and r_1 is the riskless net rate of return in units of the *numeraire*.

The consumption basket aggregates home and foreign goods. Differently from the partial equilibrium setting, I assume here a C.E.S. aggregate with elasticity of substitution different from 1. The reason is that, as shown by Cole and Obstfeld (1991) and Corsetti and Pesenti (2001), in the presence of unitary elasticity of substitution between home and foreign goods, there are no intertemporal transfers of wealth across countries (i.e., no current account movements). Therefore, the consumption basket in the Home country is defined to be:

$$C_t = \left[(C_t^s)^{\frac{\theta-1}{\theta}} + (C_t^m)^{\frac{\theta-1}{\theta}} \right]^{\frac{\theta}{1-\theta}},$$

where θ is the elasticity of substitution between goods and services, assumed to be larger than 1. C_t^s represents the consumption of home goods in Home at time t, while C_t^m is the consumption of foreign good in Home at time t. C_t^*, C_t^{s*}, and C_t^{m*} are defined in analogous fashion. The price indexes in Home and Foreign are respectively:

$$P_t = \left[(p_t^s)^{1-\theta} + (\tau_t^m)^{1-\theta} \right]^{\frac{1}{1-\theta}}, \qquad P_t^* = \left[(\tau_t^s p_t^s)^{1-\theta} + 1 \right]^{\frac{1}{1-\theta}}. \tag{1.16}$$

The intertemporal optimization problem yields Euler equations for both Home and Foreign that are identical to equation (1.7):

$$C_1 = \beta^{-\sigma} \left((1+r_1) \left(\frac{P_1}{P_2} \right) \right)^{-\sigma} C_2 \qquad C_1^* = \beta^{-\sigma} \left((1+r_1) \left(\frac{P_1^*}{P_2^*} \right) \right)^{-\sigma} C_2^*. \tag{1.17}$$

The intratemporal optimization decisions give the following demand equations for $t = \{1,2\}$:

$$C_t^s = \left(\frac{p_t^s}{P_t} \right)^{-\theta} C_t, \qquad C_t^{s*} = \left(\frac{\tau_s^h p_t^s}{P_t^*} \right)^{-\theta} C_t^*, \tag{1.18}$$

$$C_t^m = \left(\frac{\tau_t^m}{P_1} \right)^{-\theta} C_t, \qquad C_t^{m*} = \left(\frac{1}{P_t^*} \right)^{-\theta} C_t^*. \tag{1.19}$$

To close the model, we must impose goods and bond market clearing conditions. The nature of the iceberg trade costs implies the following goods market clearing conditions:

$$Y_t^s = C_t^s + \tau_t^s C_t^{s*}, \tag{1.20}$$

$$Y_t^{m*} = \tau_t^m C_t^m + C_t^{m*}. \tag{1.21}$$

Finally, bond market clearing requires:

$$B_1 + B_1^* = 0. \tag{1.22}$$

We thus have 21 endogenous variables $(C_t, C_t^*, P_t, P_t^*, C_t^s, C_t^{s*}, C_t^m, C_t^{m*}, p_t^s, B_1, B_1^*, r_1)$ with $t = \{1, 2\}$. The 21 equations (1.13)-(1.19) and (1.21), together with the evolution of the exogenous variables Y_t^j and τ_t^j (with $t = \{1, 2\}$ and $j = h, f$) completely characterize the equilibrium of this economy.[26]

Unfortunately, one cannot obtain closed-form solutions without unitary elasticity of substitution between home and foreign goods. To make the results transparent, instead of relying on numerical examples, I will present analytical results based on the log-linearized version of the model around a symmetric steady state.

1.5.1 A Symmetric Steady State

The analysis below is based on a log-linearization of the model around a symmetric steady state where $\bar{p^s} = \bar{p^{m*}} = 1$, $\bar{B_1} = \bar{B_1^*} = 0$, $\bar{Y^s} = \bar{Y^{m*}} = \bar{Y}$, and $\bar{\tau^s} = \bar{\tau^m} = \tau$.

[26] By Walras Law', the clearing of the bond and service markets implies the clearing of the manufacturing market.

In this symmetric steady state, price indexes are equal:

$$\bar{P} = \bar{P}^{*} = \left(1 + \tau^{1-\theta}\right)^{\frac{1}{1-\theta}}. \tag{1.23}$$

Moreover, we have:

$$\bar{C} = \bar{C*} = \frac{\bar{Y}}{\bar{P}}, \tag{1.24}$$

$$\bar{C^s} = \bar{C^{m*}} = \bar{P}^{\theta}\bar{C}, \tag{1.25}$$

$$\bar{C^m} = \bar{C^{s*}} = \tau^{-\theta}\bar{P}^{\theta}\bar{C}. \tag{1.26}$$

Finally the Home share of consumption of the home good is equal to the Foreign share of consumption of the foreign good:

$$\frac{\bar{C^s}}{\bar{C^s} + \tau\bar{C^{s*}}} = \frac{\bar{C^{m*}}}{\tau\bar{C^m} + \bar{C^{m*}}} = s_h = \frac{1}{1+\tau^{1-\theta}}. \tag{1.27}$$

Notice that the foreign share of consumption of the home good includes also the amounts lost to trade costs. On the other hand, the Home share of consumption of the foreign good is:

$$\frac{\tau\bar{C^m}}{\tau\bar{C^m} + \bar{C^{m*}}} = \frac{\tau\bar{C^{s*}}}{\bar{C^s} + \tau\bar{C^{s*}}} = s_f = \frac{\tau^{1-\theta}}{1+\tau^{1-\theta}} \tag{1.28}$$

Consistent with intuition, it is straightforward to check that $\frac{\partial s_h}{\partial \tau} > 0$ and $\frac{\partial s_f}{\partial \tau} < 0$. In other words, the introduction of the trade costs creates home bias in this setting even in absence of home bias in preferences.[27] Finally, symmetry implies that $s_h = 1 - s_f$. This property is extremely useful in the process of log-linearization.

1.5.2 The Log-Linear Model

I denote percentage deviations from the symmetric steady state with a hat. So $\hat{x} = log\left(\frac{x}{\bar{x}}\right)$, where $\bar{x}$ is the value of x at the symmetric steady state. The details of the log-linerarization and the solution of the model are described in the appendix. To focus my attention on the effect of trade costs, from now on I assume that endowments are constant. The Euler equations take the log-linear form:

$$\hat{C}_1 = -\sigma(1-\beta)\hat{r_1} - \sigma\hat{P}_1 + \sigma\hat{P}_2 + \hat{C}_2, \tag{1.29}$$

$$\hat{C}_1^* = -\sigma(1-\beta)\hat{r_1} - \sigma\hat{P}_1^* + \sigma\hat{P}_2^* + \hat{C}_2^*. \tag{1.30}$$

The log-linear versions of the period-1 budget constraint in Home and Foreign are:

$$\hat{B}_1 = \hat{p_1^s} - \hat{P}_1 - \hat{C}_1, \tag{1.31}$$

$$\hat{B}_1^* = -\hat{P}_1^* - \hat{C}_1^*, \tag{1.32}$$

[27]A point already made by Obstfeld and Rogoff (2001).

where importantly the current account the percentage deviation from the equilibrium output $\bar{Y}$.[28]. The budget constraints for period 2 are:

$$\hat{C}_2 = \hat{p_2^s} - \hat{P}_2 + \frac{1}{\beta}\hat{B}_1, \tag{1.33}$$

$$\hat{C}_2^* = -\hat{P}_2^* + \frac{1}{\beta}\hat{B}_1^*. \tag{1.34}$$

Taking the difference between (1.31) and (1.32) and imposing the bond market clearing condition, we get the following expression for the current account of the Home country in period 1 (equivalent to the country's net foreign asset at the end of the period):

$$2\hat{B}_1 = \hat{p_1^s} - \left(\hat{P}_1 - \hat{P}_1^*\right) - \left(\hat{C}_1 - \hat{C}_1^*\right). \tag{1.35}$$

Equation (1.35) expresses the current account of the Home country as a function of the terms of trade (p_1^s), the real exchange rate and the consumption differential. Everything else equal, an improvement of the terms of trade would lead to a current account surplus and a real appreciation to a current account deficit. An increased consumption differential between the Home and the Foreign country would lead to a current account deficit at Home. Using the difference between (1.29) and (1.30) and the difference between (1.33) and (1.34), we can rewrite (1.35) as

$$\frac{2(1+\beta)}{\beta}\hat{B}_1 = \hat{p_1^s} - \hat{p_2^s} + (\sigma - 1)\left(\hat{P}_1 - \hat{P}_2\right) - (\sigma - 1)\left(\hat{P}_1^* - \hat{P}_2^*\right). \tag{1.36}$$

Equation (1.36) allows us to interpret the evolution of Home's current account as depend-

[28]This is necessary because net foreign asset are zero in the symmetric steady state

ing on four factors. The first two represent a wealth effect. All else equal, consumption smoothing tends to push the Home current account toward surplus (deficit) in case of an increase of the value of the home endowment relative to the foreign endowment in period 1 (period 2). The next two terms represent a substitution effect. All else equal, if the intertemporal elasticity of substitution is larger than 1, an increase of the home price index in period 2 relative to period 1 tends to push Home's current account toward deficit, as would a decrease in the foreign price index in period 2 relative to period 1. These channels are the analogues to those explored in the partial equilibrium model of the previous section.

Obviously, one must solve fully the model to have the impact of the different exogenous variables on the current account. While the appendix explains the procedure in detail, I will give only a quick sketch here. For both periods, I substitute the budget constraints into the demand functions for services, and then I use the goods market clearing conditions to solve for $\hat{p_t^s}$ as function of the trade costs and $\hat{B}_1$ (imposing bonds market condition eliminates $\hat{B}_1^*$ from the system). I then express all the four elements of equation (1.36) as functions of the trade costs and $\hat{B}_1$. Finally, I substitute these functions back into equation (1.36). This allows me to express Home's current account only as function of the exogenous trade costs:

$$\hat{B}_1 = -\eta \left(\hat{\tau_1^s} - \hat{\tau_1^m} \right) + \eta \left(\hat{\tau_2^s} - \hat{\tau_2^m} \right) \tag{1.37}$$

where η is a function of the structural parameters of the model $(\beta, \theta, \sigma, \tau)$. η is a positive number as long as $\theta > 1$ and the elasticity of intertemporal substitution is sufficiently large.[29]

[29]The requirement here is weaker than $\sigma > 1$. More precisely, it is sufficient that $\sigma > 1 - g(\beta, \theta, \tau)$. See

Equation (1.37) is the key equation. It is important to notice that the relevant shock is the change in the trade cost the home good (services) *relative* to the change of the transport cost of the foreign good (manufacturing). Any symmetric trade liberalization in which the trade costs for manufacturing and services move in the same way would not have any impact on the current account. On the other hand, asymmetric trade liberalization processes for which $\left(\hat{\tau_1^s} - \hat{\tau_1^m}\right) > 0$ and/or $\left(\hat{\tau_2^s} - \hat{\tau_2^m}\right) < 0$ push the current account of the Home country into deficit.

More generally, equation (1.37) challenges the view that trade policies cannot influence the trade balance because they cannot affect savings and investment decisions.[30] While this is certainly true in static settings, things can be different in dynamic settings where the timing of the trade policy potentially matters for saving and investment (which are intertemporal decisions).[31]

1.6 Quantitative Relevance

This section evaluates the quantitative relevance of the explanation for global imbalances proposed in this paper I begin by transforming the two-period model presented in Section 4 into a multi-period model by letting the time horizon of the model become infinite and replacing the flow budget constraints (1.14) and (1.15) with the following budget constraint:[32]

the appendix for details.

[30] see for instance Lamy (2010).

[31] Obviously here the point is made only for savings.

[32] (1.38) is the budget constraint for Home. A similar constraint hold abroad.

$$B_t + \frac{\delta}{2}(B_t)^2 = (1 + r_{t-1})B_{t-1} + p_t^s Y_t^s - P_t C_t + T, \quad (1.38)$$

where $\frac{\delta}{2}(B_t)^2$ is an adjustment cost for bond holding that makes the model stationary (following Turnovsky, 1985). The adjustment cost is rebated in lump-sum fashion to households (so in equilibrium $T_t = \frac{\delta}{2}(B_t)^2$).[33]

The exercise I perform consists of solving the model in levels under the assumption of perfect foresight and using the profile of reductions of the trade costs found in Section 1.3 as the exogenous driving force.

Table 1.5 reports the calibration that I use in the experiment. I set β to 0.96 to have a model at annual frequency. I set σ 1 and θ 1.5 (following Backus, Kehoe and Kydland, 1994). I set τ to get us a ratio of imports to GDP of 15%, implying $\tau = 2.7$. Although this might seem an extreme value for τ, it is exactly the value suggested by Anderson and Van Wincoop (2004). Importantly, notice that the value of the trade cost is calibrated in this exercise. It is only the profile of the reduction in the trade costs that is taken from the empirical evidence in Section 2.

I proceed first considering a symmetric reduction in the trade costs in both sectors equal to that observed for manufacturing for the period 1994-2005. Then I consider a situation where the trade cost in manufacturing follows the pattern I found in section 1.3 for the years 1994 to 2005, while the trade cost for services is fixed for the first 12 years and then follows the same pattern of decrease as the trade cost in manufacturing for the following 12 years.[34]

[33]The calibrated value for the parameter δ is 0.00025, to ensure that besides delivering stationarity, the bond adjustment cost does not interfere with any of the economically relevant dynamics of the model.

[34]Obviously this is a strong assumption, but it is a reasonable way of capturing the evidence.

Figure 1.15 reports the result for key variables of the model in the case of a symmetric reduction in transport costs. Consistent with intuition, and with equation (1.37), there is no change in either the net foreign asset position or the net exports of either country. Consumption increases in both countries because of the reduction in the price indexes, and the only relative effect is a change in the allocation of consumption in both countries toward the good produced by the other country.

Figure 1.16 reports the results for the asymmetric liberalization process. The home country (whose goods are "liberalized" later) accumulates net foreign liabilities in anticipation of a future improvement of the relative price of its own good (an anticipated future improvement of its terms of trade). Net exports first decrease and then increase, when the home country experiences the export boom that follows the liberalization of trade in its own good. Consumption at Home initially increases and then starts declining slightly, while the consumption in Foreign country declines slightly at the beginning and starts increasing later, driven by the import boom generated by the reduction in the trade costs for the home good.

In order to gauge the quantitative relevance of this phenomenon, Figure 1.17 plots the results obtained for Home's current account under the symmetric and asymmetric liberalization episodes and the current account deficit of the U.S. for the same period (both scaled by GDP). As the figure shows, with the asymmetric trade liberalization process, the model generates a current account deficit of about 1% of GDP for 2006 (when the U.S. deficit peaked), equal to roughly 20% of the data counterpart.

In Figure 1.18, I repeat the same exercise using different values for the intertemporal elasticity of substitution. Pushing σ up to 2 (the value found in Gruber, 2005, and recently proposed by Barro, 2009) the model generates a current account deficit to GDP ratio of 3.5%. Lowering σ to 0.5, the model is unable to generate a current account deficit.

While the exact extent of the current account deficit produced by the model is certainly sensitive to parameter values, I conclude that the story presented in this paper is potentially relevant not only qualitatively, but also quantitatively. Obviously, a more complex and realistic model would be necessary to evaluate the quantitative relevance of this explanation for global imbalances more carefully. I leave this for future research.

1.7 Conclusion

This paper has proposed a new explanation for the emergence of global imbalances. The focus is on the interplay between the comparative advantage of the U.S. in services and the asymmetric trade liberalization process in goods trade versus service trade that took place in the last two decades, particularly since the mid-nineties. This explanation complements existing ones by adding a quantitative non-negligible "real" piece to the global imbalances puzzle.

The first and most obvious policy implication of this paper is that liberalization of trade in services represents a possible margin of adjustment that might alleviate global imbalances without necessarily implying a large depreciation of the dollar.[35] A similar suggestion has recently been proposed by Claessens, Evenett, and Hoekman (2010) and Kowalski and Lesher (2010).

The second, indirect, policy implication is that, if one believes the asymmetry in the

[35] As suggested by the work of Blanchard, Giavazzi, and Sa (2005) and Obstfeld and Rogoff (2005).

liberalization of service trade and manufacturing trade will continue[36], then there could be a case for a "return" of the U.S. to manufacturing.

There are several directions for future research. First, it would be interesting to study service trade at a more disaggregated level, thus taking better account of the heterogeneity between service industries.

Second, it would be interesting to study FDI in services both empirically and theoretically to evaluate how introducing FDI would affect the conclusions of this paper. The key constraint is the lack of data on bilateral foreign affiliate sales in the service sector.

Third, it would be important to develop a more complete model that might assess more carefully the quantitative relevance of the mechanisms proposed in this paper. I plan to pursue these avenues for research in future work.

[36]Contrary to what assumed in the quantitative exercise performed in this paper.

1.8 Appendix

1.8.1 Data

The data sources used in this paper are several. The data used for Figures 1-10 and section 2 are taken from the World Bank World Development Indicators (WDI), namely the series for the GDP, the export and imports of goods and services and the income balance, the real gdp per capita, the gdp per capita growth, the private credit over GDP. The sample of countries includes the 30 OECD countries and the 48 Upper-Middle income countries according to the World Bank classification, plus China, India, Indonesia, Singapore and Saudi Arabia. The time period is 1970-2010.

In the empirical analysis of section 3 the data on trade in services come from the Trade in Service Database, developed by Francois et al (2009) using OECD, Eurostat and IMF data. The data for trade in manufacturing are taken from the UN-Comtrade database. The distance data is the population-weighted data of the distance dataset of the CEPII, while the data on common borders, language, legal origin and colonial origin are taken from Baranga (2009). Importantly, in the regressions enter also the "internal trade" (the *ii* elements), defined as production minus exports. The sample of countries includes 23 OECD countries plus China (Austria, Canada, China, Czech Republic, Germany, Denmark, Spain, Finland, France, Greece, Hungary, Iceland, Italy, Japan, Korea, Netherlands, Norway, New Zealand, Poland, Portugal, Slovakia, Sweden, UK, US). The sample is reduced because of the need for data on gross output at the sectoral level, available from the OECD-STAN database only for few countries. The output data at the sectoral level for China are available from the *China Statistical Yearbook* I-O tables, but only for 1997,2000,2002,2005. The data for the remaining years are imputed by multiplying the value added data (GDP at sector level, available from the national income account) by the output/value added ratio at the sectoral level found the nearest available year (so using the 1997 data for the period 1994-1998, the 2000 data for the period 1999-2001, the 2002 data for the period 2002-2003 and the 2005 data for the period 2004-2005).

1.8.2 Model

The Two-Country Model

Here are the 23 equations of the model:

$$C_2 = \beta^\sigma \left((1+r_t) \left(\frac{P_1}{P_2} \right) \right)^\sigma C_1 \qquad C_2^* = \beta^\sigma \left((1+r_t) \left(\frac{P_1^*}{P_2^*} \right) \right)^\sigma C_1^* \tag{1.39}$$

$$B_1 = p_1^s Y_1^s - P_1 C_1 \qquad B_1^* = p_1^{m*} Y_1^m - P_1^* C_1^* \tag{1.40}$$

$$P_2 C_2 = p_2^s Y_2^s + (1+r_t) B_1 \qquad P_2^* C_2^* = p_2^{m*} Y_2^f + (1+r_t) B_1^* \tag{1.41}$$

$$P_t = \left[(p_t^s)^{1-\theta} + (\tau_t^m p_t^{m*})^{1-\theta} \right]^{\frac{1}{1-\theta}} \qquad P_t^* = \left[(\tau_t^s p_t^s)^{1-\theta} + (p_t^{m*})^{1-\theta} \right]^{\frac{1}{1-\theta}} \tag{1.42}$$

$$C_t^s = \left(\frac{p_t^s}{P_t}\right)^{-\theta} C_t \qquad C_t^{s*} = \left(\frac{\tau_t^s p_t^s}{P_t^*}\right)^{-\theta} C_t^* \tag{1.43}$$

$$C_t^m = \left(\frac{\tau_t^m p_t^{m*}}{P_t}\right)^{-\theta} C_t \qquad C_t^{m*} = \left(\frac{p_t^{m*}}{P_t^*}\right)^{-\theta} C_t^* \tag{1.44}$$

$$Y_t^s = C_t^s + \tau_t^s C_t^{s*} \tag{1.45}$$

$$Y_t^{m*} = \tau_t^m C_t^m + C_t^{m*} \tag{1.46}$$

$$B_t + B_t^* = 0 \tag{1.47}$$

So we have 23 endogenous variables $(C_t, C_t^*, P_t, P_t^*, C_t^s, C_t^{s*}, C_t^m, C_t^{m*}, p_t^s, B_1, B_1^*, r_1)$ with $t = [1, 2]$. The 23 equation (1.39)-(1.47), together with the evolution of the exogenous variables Y_t^j and τ_t^j (with $t = [1, 2]$ and $j = [s, m]$) completely characterize the equilibrium of this economy. I fix p_t^{m*} to be the *numeraire*. Finally, by Walras Law, I can eliminate the manufacturing goods market clearing conditions.

The Complete Log Linearized Model

I denote with a $\hat{}$ the percentage deviations from the symmetric steady state. So $\hat{x} = log\left(\frac{x}{x^*}\right)$, where x^* is the value of x at the symmetric equilibrium. Log-linearizing the model around the symmetric steady state described in the main texts gives us:

$$\hat{C}_2 = \sigma(1-\beta)\hat{r_1} + \sigma\hat{P}_1 - \sigma\hat{P}_2 + \hat{C}_1 \qquad \hat{C}_2^* = \sigma(1-\beta)\hat{r_1} + \sigma\hat{P}_1^* - \sigma\hat{P}_2^* + \hat{C}_1^* \tag{1.48}$$

$$\hat{B}_1 = \hat{p_1^s} - \hat{P}_1 - \hat{C}_1 \qquad \hat{B}_1^* = -\hat{P}_1^* - \hat{C}_1^* \tag{1.49}$$

$$\hat{C}_2 = \hat{p_2^s} - \hat{P}_2 + \frac{1}{\beta}\hat{B}_1 \qquad \hat{C}_2^* = -\hat{P}_2^* + \frac{1}{\beta}\hat{B}_1^* \tag{1.50}$$

$$\hat{C_t^s} = -\theta\left(\hat{p_t^s} - \hat{P}_t\right) + \hat{C}_t \qquad \hat{C_t^{s*}} = -\theta\left(\hat{p_t^s} + \hat{\tau_t^s} - \hat{P}_t^*\right) + \hat{C}_t^* \tag{1.51}$$

$$\hat{C_t^m} = -\theta\left(+\hat{\tau_t^m} - \hat{P}_t\right) + \hat{C}_t \qquad \hat{C_t^{m*}} = \theta\hat{P}_t^* + \hat{C}_t^* \tag{1.52}$$

$$\hat{P}_t = s_h\hat{p_t^s} + s_f\left(\hat{\tau_t^m}\right) \qquad \hat{P}_t^* = (1-s_h)\left(\hat{p_t^s} + \hat{\tau_t^s}\right) \tag{1.53}$$

$$s_h\hat{C_t^s} + (1-s_h)\left(\hat{\tau_t^s} + \hat{C_t^{s*}}\right) = 0 \tag{1.54}$$

$$\hat{B}_1 + \hat{B}_1^* = 0 \tag{1.55}$$

Solution

In order to solve the model, I plug into equation (1.54) the home and foreign version of equation (1.51) for period one. I then substitute in the resulting equation the Price indexes and the aggregate consumption levels using the period 1 budget constraints (1.49) and the Price index definitions (1.53). This gives me an equation in two unknowns, from which I derive an expression for $\hat{p_1^s}$ as function of $\hat{\tau_1^s}$, $\hat{\tau_1^m}$ and $\hat{B}_1$:

$$\hat{p_1^s} = \gamma_1 \left(\hat{\tau_1^s} - \hat{\tau_1^m}\right) - \beta\gamma_0 \hat{B}_1 \tag{1.56}$$

Where I defined the following parameters (some of the signs are valid only under the restriction $\theta > 1$):

$$\begin{aligned}
\alpha_0 &= \frac{s_h - s_f}{\beta} > 0 \\
\alpha_1 &= s_f s_h(\theta - 1) > 0 \\
\alpha_2 &= 2\alpha_1 + s_f > 0 \\
\gamma_0 &= \frac{\alpha_0}{\alpha_2} > 0 \\
\gamma_1 &= -\frac{\alpha_1}{\alpha_2} < 0
\end{aligned} \tag{1.57}$$

Moreover, It is easy to show how:

$$\hat{P}_1 - \hat{P}_1^* = \gamma_2 \left(\hat{\tau_1^s} - \hat{\tau_1^m}\right) - (s_h - s_f)\beta\gamma_0 \hat{B}_1 \tag{1.58}$$

with $\gamma_2 = (s_h - s_f)\gamma_1 - s_f < 0$.
Repeating the same procedure for period two, I get a very similar expression:

$$\hat{p_2^s} = \gamma_1 \left(\hat{\tau_2^s} - \hat{\tau_2^m}\right) + \gamma_0 \hat{B}_1 \tag{1.59}$$

and

$$\hat{P}_2 - \hat{P}_2^* = \gamma_2 \left(\hat{\tau_2^s} - \hat{\tau_2^m}\right) + (s_h - s_f)\gamma_0 \hat{B}_1 \tag{1.60}$$

Plugging back equations (1.56)-(1.60) into equation (1.36) after rearranging and defining

$$\eta = -\frac{\beta\left(\gamma_1 + (\sigma - 1)\gamma_2\right)}{2(1+\beta) + \beta\gamma_0(1+\beta)\left[1 + (s_h - s_f)(\sigma - 1)\right]} > 0 \tag{1.61}$$

gives equation (1.37) in the main text. From Equation (1.61) is possible to derive the restriction on the intertemporal elasticity of substitution that makes η a positive number (given $\theta > 1$). In particular, it has to be $\sigma > 1 - \frac{\frac{\alpha_1}{\alpha_2}}{(s_h - s_f)\frac{\alpha_1}{\alpha_2} + s_f}$

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Table 1.1: **Cross Section Analysis. Dependent Variable: Current Account/GDP**

Year	2005	2005	2005	1995	1995	1995
PAST RCA_SERV	**-10.147*** **	-10.162***	-9.757***	**-1.847** **	-2.391***	-2.500***
	(1.190)	(1.192)	(1.062)	(0.748)	(0.795)	(0.750)
GDP_PC_GROWTH		0.296	0.178		-0.208	-0.162
		(0.275)	(0.245)		(0.211)	(0.200)
GDP_PC_REAL		0.000	0.001***		0.000***	0.000***
		(0.000)	(0.000)		(0.000)	(0.000)
OPENNESS		3.229	2.609		2.778**	3.417***
		(1.982)	(1.771)		(1.331)	(1.274)
CREDIT			-0.106***			-0.063***
			(0.026)			(0.022)
R-squared	**0.506**	0.526	0.623	**0.077**	0.197	0.287
N	71	71	70	62	61	61

Standard Errors in Parenthesis

*,**,*** Statistically Significant at 10%, 5% and 1%

Table 1.2: **Panel Analysis. Dependent Variable: Current Account/GDP (5-years Averages**

Years	1996-2010	1996-2010	1996-2010	1970-1995	1970-1995	1970-1995
Lagged_RCA_SERV	**-5.829*** **	-5.427***	-4.829***	**-0.403**	-0.867	-0.670
	(1.807)	(1.789)	(1.824)	(1.363)	(1.344)	(1.302)
GDP_PC_GROWTH		0.901***	0.810***		0.120	-0.032
		(0.223)	(0.232)		(0.175)	(0.177)
REAL_GDP_PC		0.001	0.001*		0.001***	0.001***
		(0.000)	(0.000)		(0.000)	(0.000)
OPENNESS		2.570	2.911		2.292	4.115
		(4.219)	(4.226)		(3.542)	(3.534)
CREDIT			-0.030			-0.101***
			(0.023)			(0.027)
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	**0.42**	0.42	0.41	**0.08**	0.09	0.07
N	223	223	222	232	225	220

Standard Errors in Parenthesis

*,**,*** Statistically Significant at 10%, 5% and 1%

Table 1.3: **PPML Results, Manufacturing and Service Trade**

Industry	**Manuf**	**Manuf**	**Total Services**	**Total Services**	**Business Services**	**Business Services**
Year	**1999**	**2005**	**1999**	**2005**	**1999**	**2005**
distance	-1.009***	-1.145***	-0.830***	-0.858***	-0.826***	-0.829***
	(0.083)	(0.086)	(0.141)	(0.096)	(0.170)	(0.103)
border	0.078	0.089	0.164	0.116	0.227	0.113
	(0.244)	(0.178)	(0.475)	(0.297)	(0.503)	(0.275)
language	-0.208	-0.339**	-0.025	-0.106	-0.111	-0.064
	(0.171)	(0.138)	(0.283)	(0.175)	(0.339)	(0.159)
colonial	0.608***	0.461**	1.086***	0.826***	1.127***	0.861***
	(0.186)	(0.182)	(0.205)	(0.145)	(0.192)	(0.133)
legal	0.528***	0.340***	0.550***	0.584***	0.563***	0.567***
	(0.116)	(0.103)	(0.167)	(0.120)	(0.172)	(0.121)
smctry	2.824***	2.095***	5.755***	5.371***	4.481***	4.221***
	(0.199)	(0.195)	(0.335)	(0.230)	(0.400)	(0.237)
Imp FE	Yes	Yes	Yes	Yes	Yes	Yes
Exp FE	Yes	Yes	Yes	Yes	Yes	Yes
N	576	576	449	532	356	457

Standard Errors in Parenthesis

*,**,*** Statistically Significant at 10%, 5% and 1%

Table 1.4: **CHB, Manufacturing and Services**

Industry	**Manuf**	**Manuf**	**Manuf**	**Total Services**	**Total Services**	**Business Services**	**Business Services**
Year	**1994**	**1999**	**2005**	**1999**	**2005**	**1999**	**2005**
USA	3.2	2.9	3.3	2.3	2.4	2.1	2.2
JPN	3.8	5.1	6.3	5.9	8.0	6.5	8.6
DEU	9.5	10.0	9.7	12.2	13.3	11.0	11.3
GBR	20.2	17.9	19.2	15.3	14.0	13.0	12.0
FRA	15.0	14.6	13.2	17.0	15.6	14.8	13.0
ITA	16.5	15.1	14.0	20.7	19.0	20.6	17.9
CHN	17.8	11.1	6.0	43.4	21.4		
CAN	30.4	25.8	21.0	39.0	33.0	36.4	29.7
ESP	30.8	26.5	20.5	43.3	33.5	44.7	33.4
KOR	26.8	26.7	18.6	65.9	50.3	60.5	44.9
NLD	59.1	57.1	51.7	57.6	53.6	49.3	45.4
SWE	85.6	65.2	57.4	93.3	92.0	80.6	77.1
AUT	84.7	79.7	65.5	122.3	116.3	114.0	100.9
POL	129.3	88.8	59.1	161.6	122.0		
DNK	149.2	143.5	122.3	145.1	133.4	142.6	114.5
NOR	147.8	119.0	91.2	164.6	136.0	134.3	105.7
GRC	198.1	187.0	152.9	205.5	159.3	287.8	210.5
PRT	137.4	122.7	105.7	214.9	190.6	224.6	189.3
FIN	157.9	119.9	102.1	213.5	198.4	214.9	208.6
CZE	197.9	140.1	75.5	431.1	270.4	358.3	199.8
NZL	401.0	445.0	341.1	385.1	286.4		
HUN	283.0	207.0	120.2	558.2	336.0	537.7	296.9
SVK	448.5	313.8	152.2	1082.3	755.7	820.8	561.0
ISL	2538.7	1802.9	1334.4	3148.9	2068.0	3161.0	1851.2

Table 1.5: **Calibration**

Parameter	**Meaning**	**Baseline**
σ	Intertemp. Elast of substitution	1
β	Discount Factor	0.96
θ	Intratemp. Elast of substitution	1.5
τ	Trade Cost	2.7
δ	Bond Adj. cost	0.00025

Figure 1.1: **Composition of the US Current Account Balance (I), Percentage of GDP**

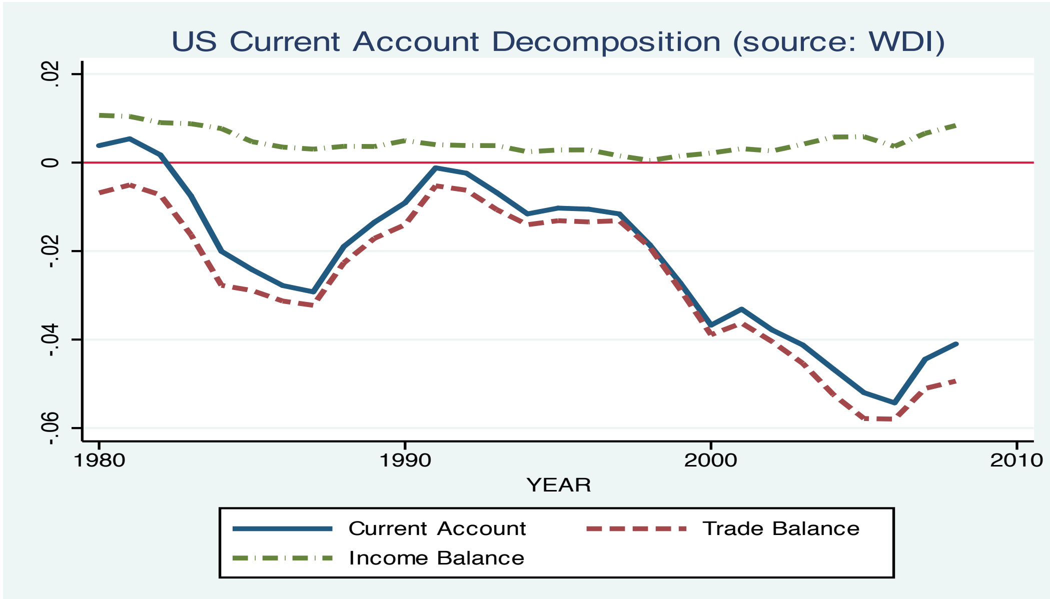


Figure 1.2: **Composition of the US Current Account Balance (II), Percentage of GDP**

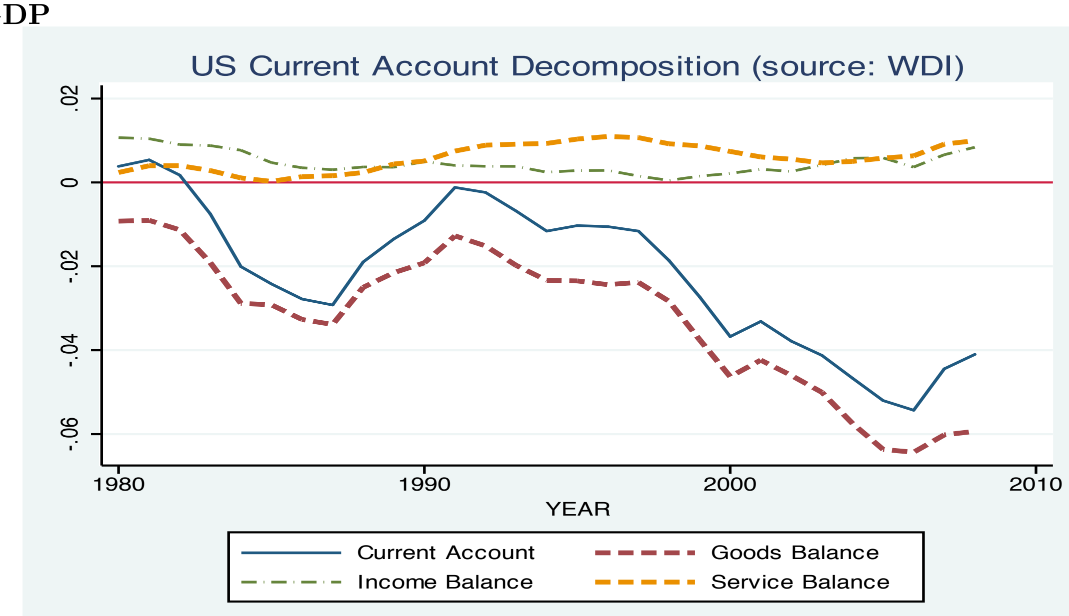


Figure 1.3: **Composition of the Japan's Current Account Balance, Percentage of GDP**

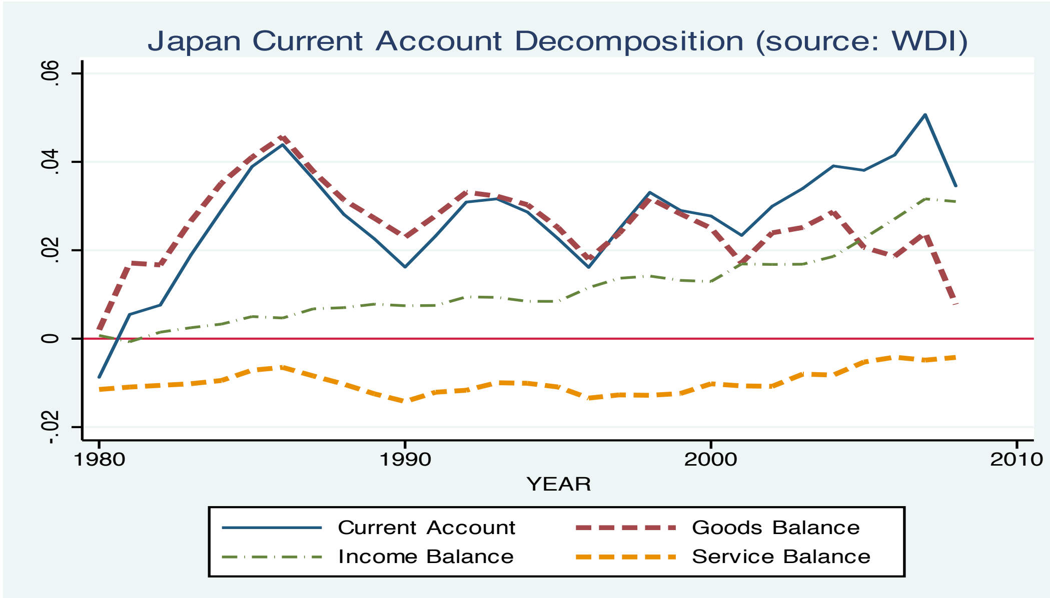


Figure 1.4: **Composition of the Germany's Current Account Balance, Percentage of GDP**

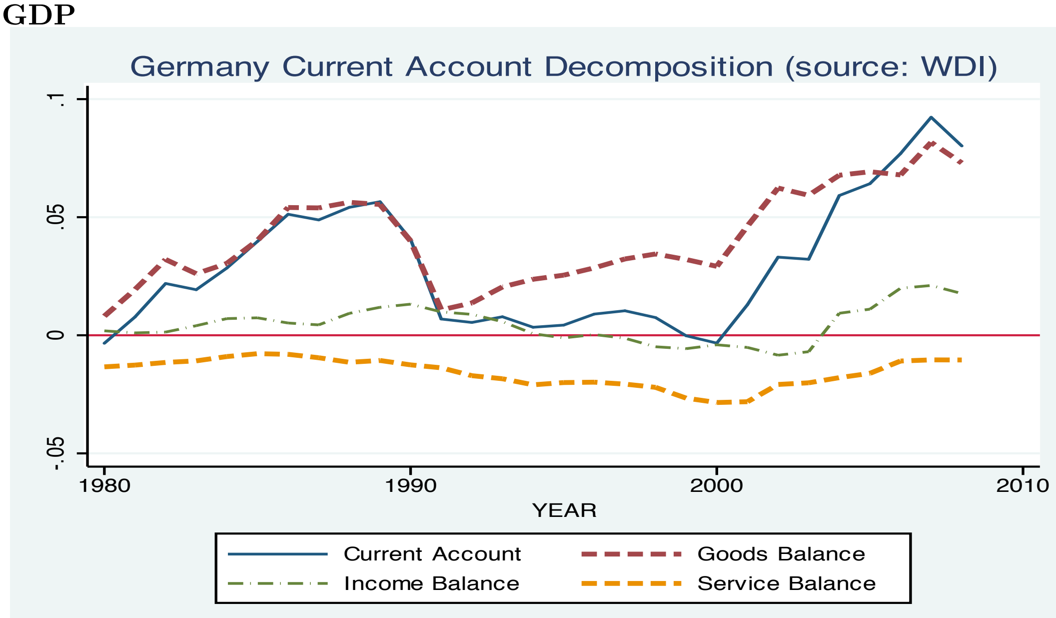


Figure 1.5: **Composition of the China's Current Account Balance, Percentage of GDP**

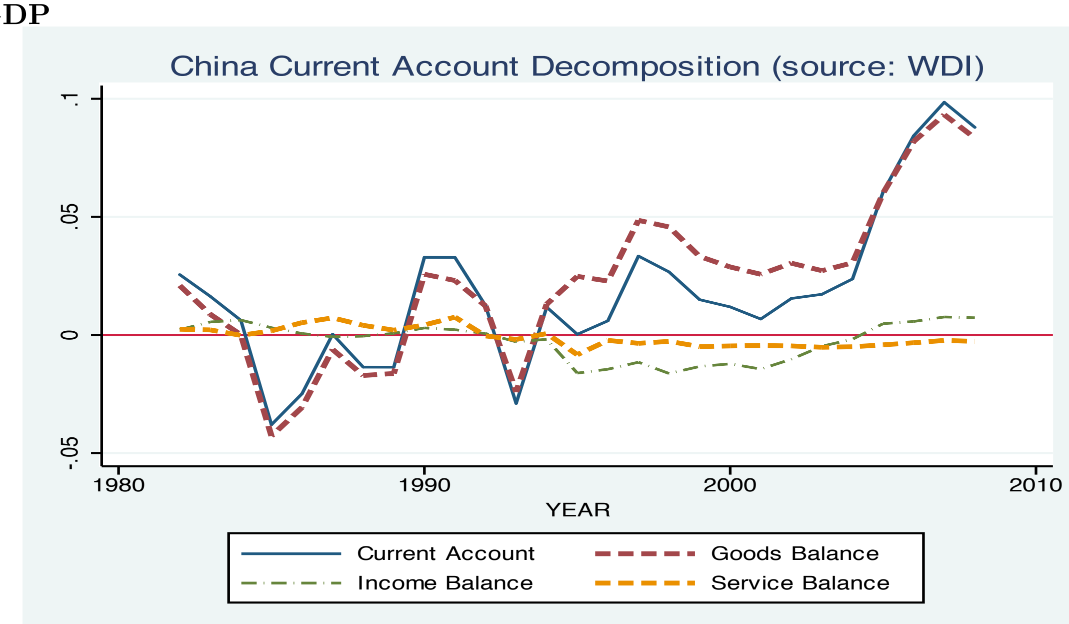


Figure 1.6: **Revealed Comparative Advantage (RCA) Index, Services, Selected Countries, 1994-2005**

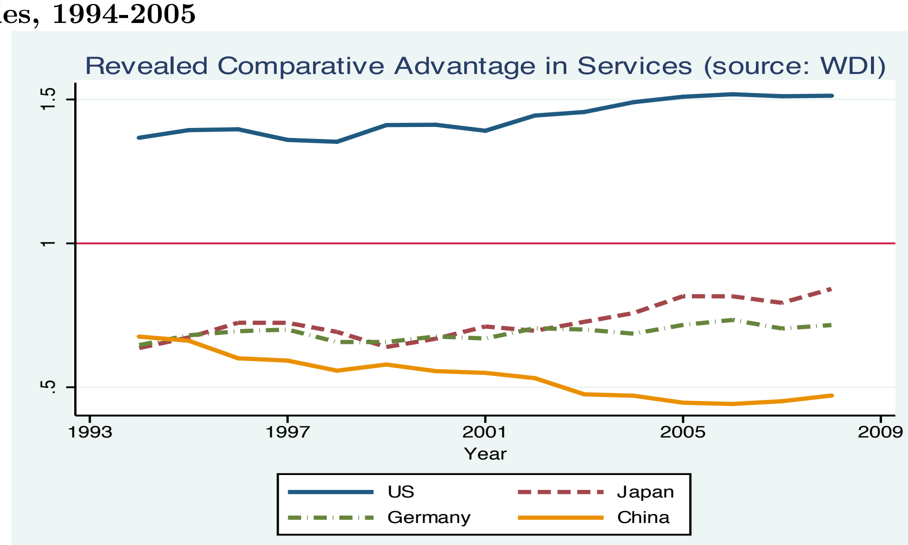


Figure 1.7: **Current Account over GDP and RCA in Services, OECD plus BRICS, 1995**

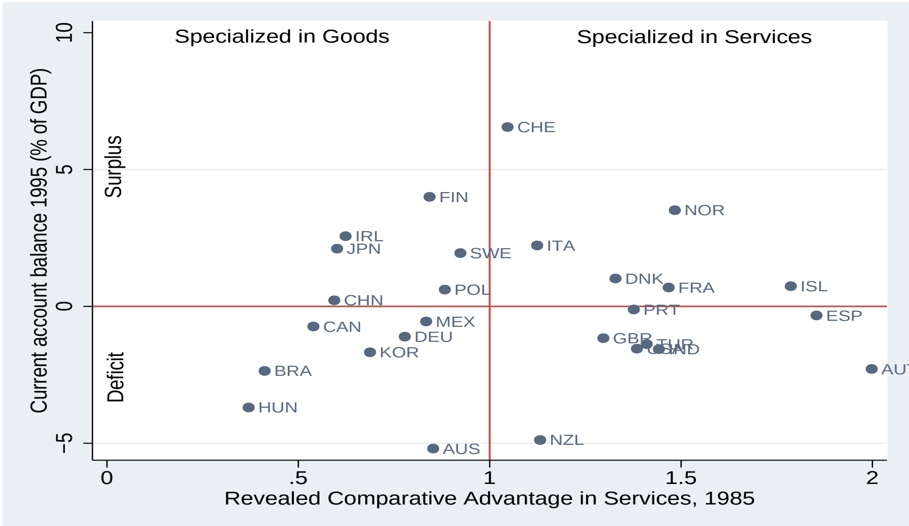


Figure 1.8: **Current Account over GDP and RCA in Services, OECD plus BRICS, 2006**

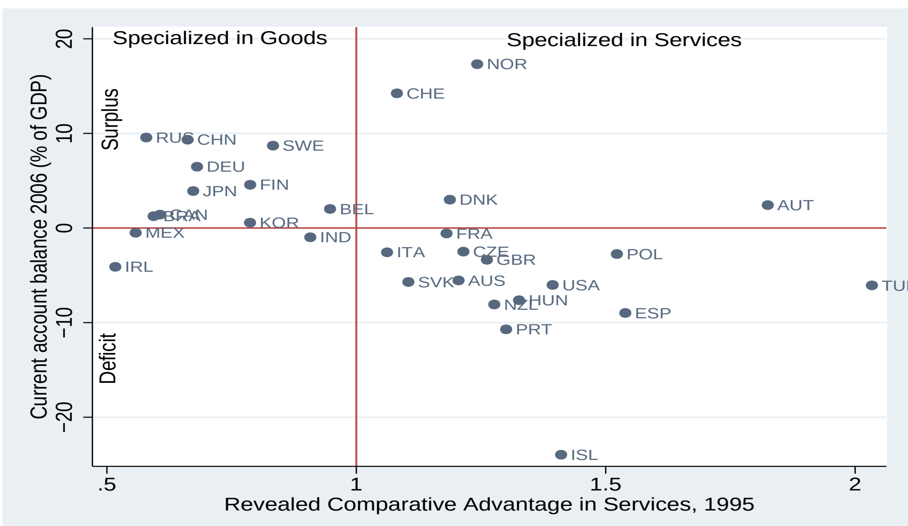


Figure 1.9: **Current Account over GDP and RCA in Services, 83 OECD and Developing Countries, 2006**

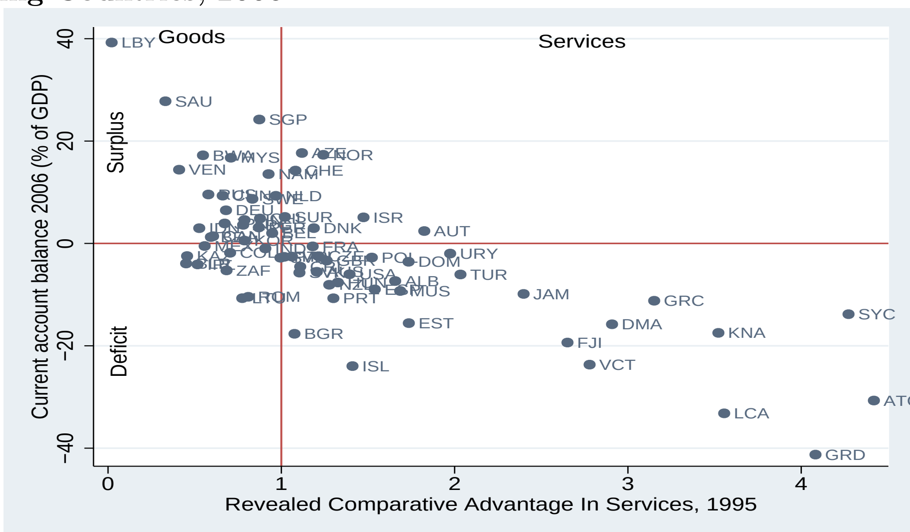


Figure 1.10: **Constructed Home Bias World Index, Manufacturing (1994-2005) and Services (1999-2005)**

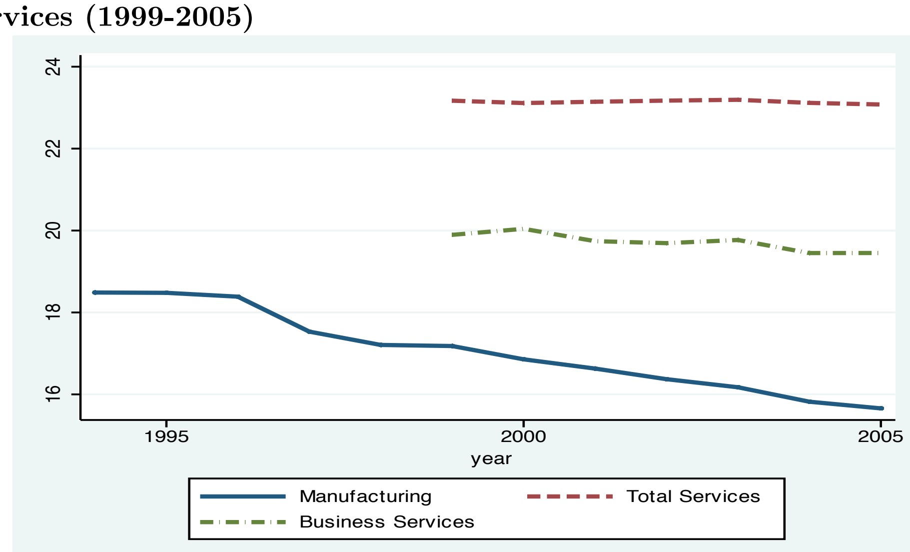


Figure 1.11: **Constructed Home Bias World Index, Manufacturing (1994-2005) and Services (1999-2005). First year=100**

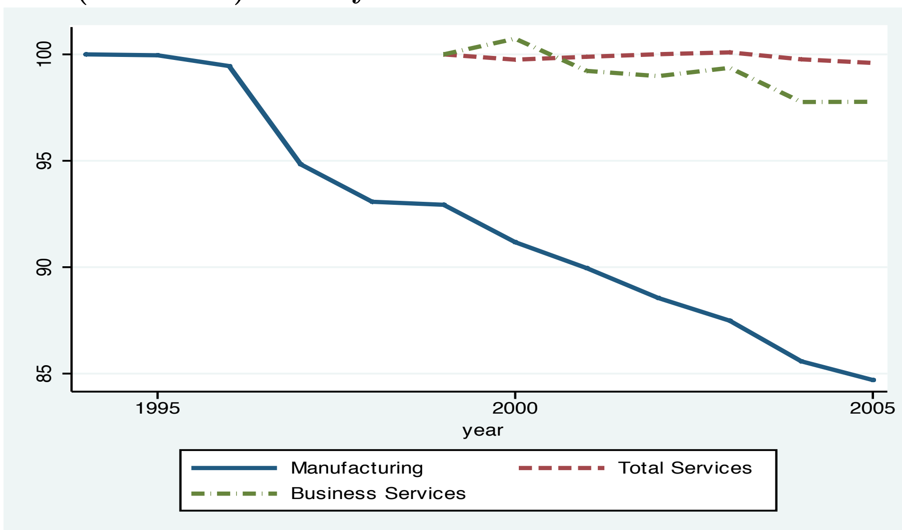


Figure 1.12: **Goods Trade vs. Service Trade: An Asymmetric Trade Liberalization Process**

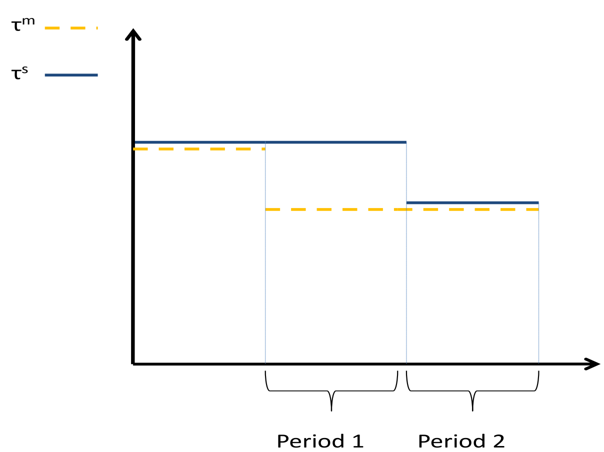


Figure 1.13: **Asymmetric Trade Liberalization and Current Account: Service Importers**

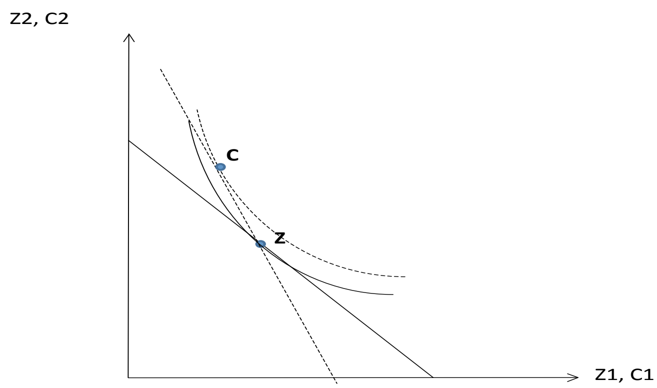


Figure 1.14: **Asymmetric Trade Liberalization and Current Account: Service Exporters**

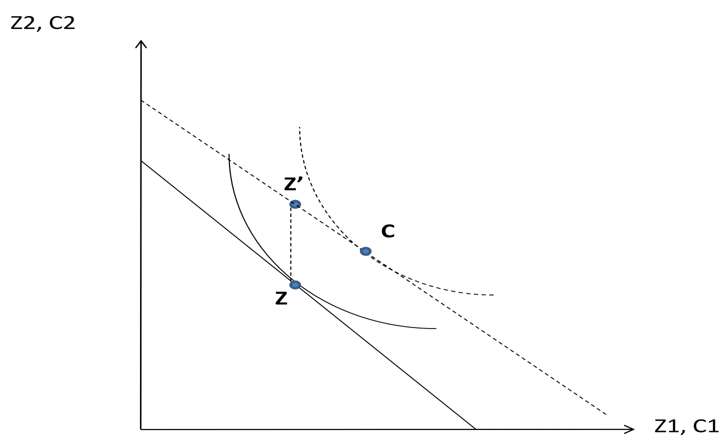


Figure 1.15: **Model Simulations: Symmetric Trade Liberalization**

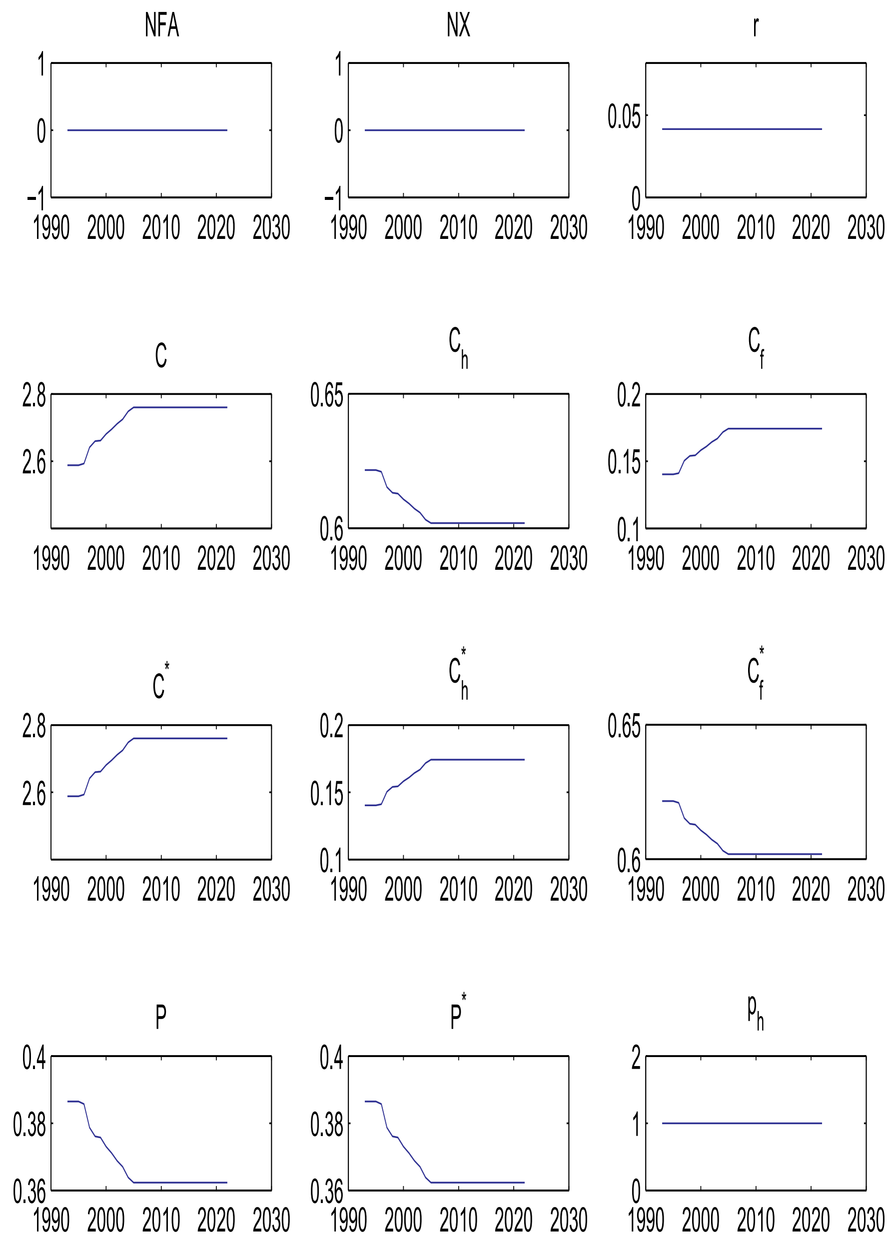


Figure 1.16: **Model Simulations: Asymmetric Trade Liberalization**

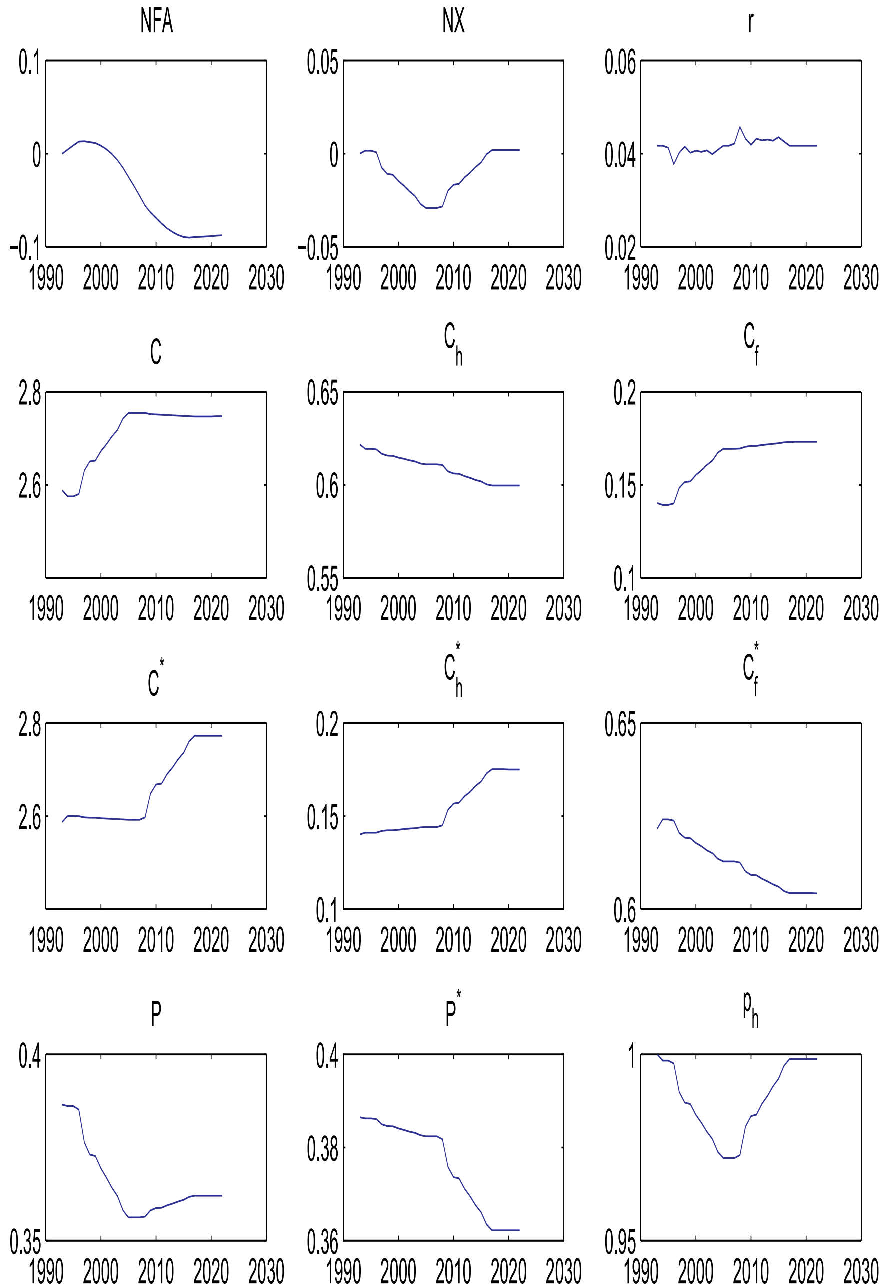


Figure 1.17: **The Quantitative Relevance: Time Series (Baseline)**

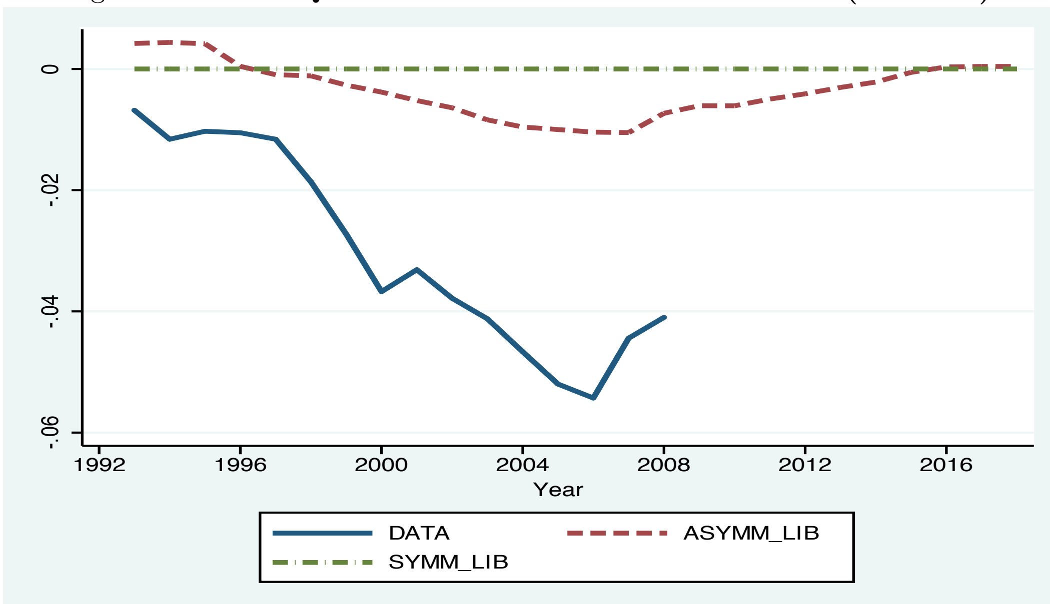
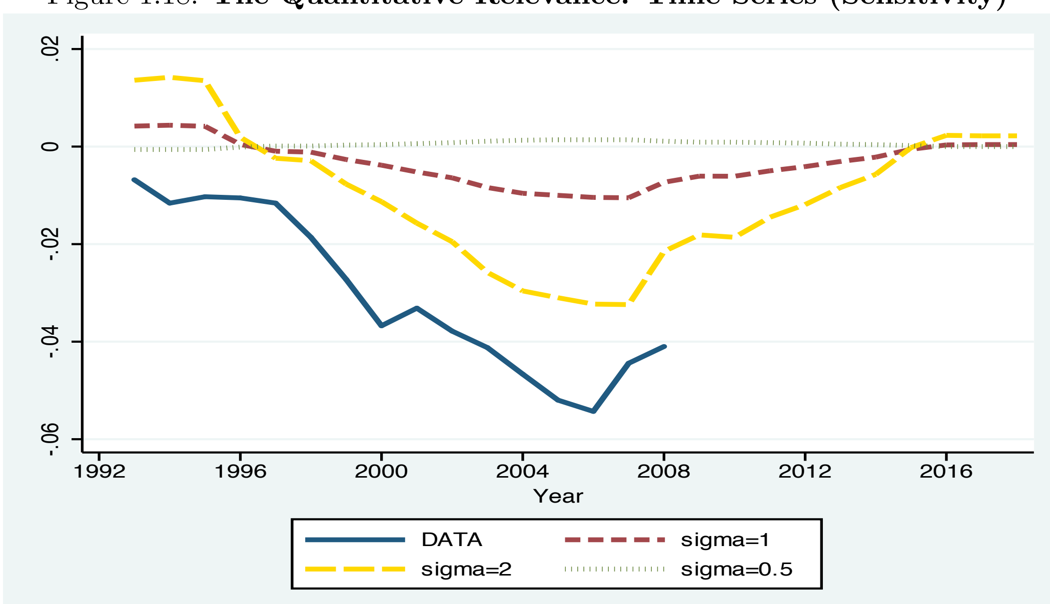


Figure 1.18: **The Quantitative Relevance: Time Series (Sensitivity)**



Chapter 2

Estimating Trade and Investment Flows: Partners and Volumes

2.1 Introduction

Three facts constitute the background of this work. First, trade and Foreign Direct Investment (FDI) have been among the fastest growing economic activities around the world in the last decades (Helpman, 2006). While clearly interconnected, these two phenomena have often been treated separately in the economics literature. An important exception is Helpman, Melitz and Yeaple (2004, henceforth HMY), who extend the Melitz (2003) model of trade to the case of trade and *horizontal* FDI.[1] Second, bilateral trade flows are characterized by the presence of a lot of zeroes (i.e. the absence of flows among many country pairs). This observation motivated Helpman, Melitz and Rubinstein (2008, henceforth HMR) to propose a two-stage estimation methodology that corrects the gravity-type specification for bilateral trade flows for selection and, more importantly, for firms' heterogeneity. Third, work by Razin and Sadka (2007) showed that selection also plays an important role in

[1]Defined as the investment abroad aimed at serving the foreign market, as opposed to *vertical* FDI, which are investments aimed at reducing costs through the vertical disintegration of the production process, such as the case of the Mexican *Maquiladoras*. Another notable exception is the work by Ramondo and Rodriguez-Claire (2009)

the FDI case, and they illustrate the advantages of using sample selection models when estimating bilateral investment flows.

In this paper, I start by showing empirical evidence from a large sample of countries for the period 2000-2007. Bilateral investment flows are almost never observed in the absence of bilateral trade flows, thus configuring an *order* of trade and investment flows. The same pattern is uncovered using firm-level data from a novel dataset (ORBIS) and then constructing a series of aggregate bilateral foreign affiliate sales (FAS) in the manufacturing sector.[2]

Consistent with this evidence, I present a model where heterogeneous firms face a proximity-concentration trade-off in deciding whether to serve foreign markets through export or FDI, along the lines of HMY. If a firm serves the foreign market through export, it pays a lower fixed cost but bears a higher variable cost due to the existence of an *iceberg* transportation cost. If it decides to invest abroad, the fixed cost is higher[3] but the variable cost is lower. Departing from HMY, I assume that investing abroad implies the existence of a *cost disadvantage* for the foreign affiliate vis a vis the domestic firms, which is conveniently defined as a fraction of the transport cost and depends on the *economic distance* between countries. Some evidence regarding this assumption is provided below. This allows me to derive the implications of the model for aggregate bilateral trade and foreign affiliate sales flows in the form of theory-based gravity-type equations.

I then suggest a two-stage estimation procedure along the lines of HMR. In the first stage, an *ordered Probit* model is used to retrieve consistent estimates of the terms needed

[2]Bilateral foreign affiliate sales are observed in the absence of trade mostly for countries like Liechenstein, Bermuda or Luxemburg, which are often only the places where the head-quarters of multinational firms are established in order to benefit from a more favorable tax-treatment of the profits.

[3]A multiple of the fixed cost of exporting.

to correct the flows equations for heterogeneity and selection. The ordered Probit is completely derived from theory and from the definition of appropriate latent variables, under the assumptions that the marginal cost in the case of investment is a fraction of the marginal cost in the case of export while the fixed cost of exporting is a multiple of the fixed cost of investing. In the second stage, maximum likelihood (ML) can be applied to the corrected trade, FDI and foreign affiliate sales flows equations.

The main results of the analysis are as follows: 1) The impact of distance, border and regional trade agreements on bilateral foreign affiliate sales becomes substantially smaller after controlling for selection and firms' heterogeneity (hence separating the impact on the extensive versus the intensive margin). 2) The same "attenuation result" is found also for the trade equations, consistently with HMR. 3) When FAS are observed, failing to take this into account when correcting for heterogeneity and selection in the trade equations leads to differences in the estimated coefficients.

This paper is linked to several strands of the literature. First, this work is related to the literature on models of trade with heterogeneous firms (Melitz, 2003 ; HMY) as well as to the gravity models of bilateral trade flows (Anderson, 1979; Anderson and van Wincoop, 2003) and to the recently proposed HMR procedure of estimating trade flows correcting for selection and heterogeneity. Santos Silva and Tenreyro (2006) provides some qualifications that will be addressed later in the paper.

Second, this work is related to the literature on FDI and FAS. Kleinert and Toubal (2009) derive gravity-type equations for bilateral FDI flows. The explanation they propose is based on the dependence of the fixed cost of exporting on distance and is different from the one proposed in this paper, based on the existence of a *cost disadvantage* for foreign firms in the purchase of local intermediate inputs. Aisbett (2007) explores the importance of Bilateral Investment Treaties on bilateral investment flows.[4] Razin and Sadka (2007) propose a detailed study of aggregate bilateral FDI flows showing the importance of selection in this context.

Third, some recent work has tried to jointly consider trade and investment flows. Aviat and Coeurdacier (2007) explore the complementarity between bilateral trade in goods and asset holdings in a simultaneous gravity equation framework. Bergstrand and Egger (2007) augment a 2x2x2 Knowledge capital model with physical capital and provide a rationale for gravity-type equations for FDI. Lastly, Lai and Zhu (2006) propose a non linear joint ML estimation for trade and foreign affiliate sales for US based multinational firms. Ramondo, Rappoport and Ruhl (2009) analyze the proximity-concentration trade-off in the presence of risk. I improve on this literature by explicitly correcting for selection and heterogeneity as in HMR.

The paper is organized as follows. Section 2.2 contains a quick glance at the data that establish the ordering of trade and investment flows. Section 2.3 contains the model and section 2.4 the empirical methodology. In section 2.5, I present the main results of the analysis and section 2.6 concludes suggesting some lines for future research.

2.2 A Glance at the Data

I will use four data sources in the paper. The first is the OECD International Investment Database, which contains information about the aggregate Outward FDI flows reported by 30 OECD countries and over 200 possible destination countries in the period 2000-2007.

[4] I'm particularly grateful to her for providing the data for preliminary work on the key idea of this paper.

Importantly, in this dataset true zeroes are distinguished from missing data, thus providing reliable information about the country pairs where no FDI takes place.

The second source of data is a firm-level dataset (ORBIS), which contains data on the location of the global owner and the level of sales for the period 2000-2006 for a sample of over 45,000 active manufacturing firms located in over 90 developed and developing countries. I use this dataset to build a series of *aggregate bilateral foreign affiliate sales*.[5]

The source for trade data and the gravity variables are more standard. I will use the UN-COMTRADE database for trade flows and the CEPII distance dataset for the gravity regressors.

I build indicator variables called TRADE, FAS and FDI to indicate the presence of positive flows. Table 2.1 reports the distribution of available observations into the four possible cases (NO TRADE-NO FDI, TRADE-NO FDI, TRADE-FDI and NO TRADE-FDI) using the OECD data for FDI and the COMTRADE data for exports. Two observations stand out. First, the number of zeroes is clearly not irrelevant both for the trade and the investment flows. The number of zero trade flows is smaller than that documented in HMR (2008) because the table excludes all the observations which report a missing for the investment flows. Second, probably most interestingly, the case of FDI- NO TRADE is very infrequent. In particular, out of 351 observations, in 66 cases the importing (host) country is Liechenstein and in 76 cases the exporting (home) country is Luxemburg. Almost half of the cases in which bilateral FDI flows are observed in the absence of trade flows are likely due to tax-treatment of corporate profits.

[5]The zeroes at the firm-level come from cases in which active firms are present in country i but do not report sales. If this is true for all the firms owned by a parent company based in country j, then the aggregate FAS from country j to country i is considered a zero.

Table 2.2 reports similar statistics obtained using the bilateral FAS dataset.[6] I selected only manufacturing firms from the ORBIS dataset. For consistency, I then retrieved from the UN-COMTRADE database the aggregate trade in manufacturing sector. The evidence is similar to that reported in Table 2.1. The number of cases in which bilateral foreign affiliate sales are observed in the absence of bilateral trade flows is infrequent. Also in this case, most of the observations in the bin NO TRADE-FAS are accounted for by fiscal paradises like Bermuda, Antilles or Cyprus.[7]

The evidence proposed in Table 2.1 and 2.2 suggests a sort of *ordering* of trade and investment flows, for which the existence of bilateral trade flows is a necessary condition for the existence of bilateral investment flows and foreign affiliate sales. The theoretical model presented in the next section implies exactly this pattern for the aggregate flows.

2.3 Theory

Consider a world economy made up of J countries. In each country, a representative consumer derives utility from a continuum of goods, defined as follows for a generic country j:

$$u_j = \left(\int_{l \in \Omega_j} x_j(l)^{\alpha} dl \right)^{\frac{1}{\alpha}} \tag{2.1}$$

where $x_j(l)$ is the consumption of product l and Ω_j is the set of available varieties in country j and $\epsilon = \frac{1}{1-\alpha} > 1$ is the elasticity of substitution, assumed to be equal across

[6] Now the four possible cases become NO TRADE-NO FAS, TRADE-NO FAS, TRADE-FAS and NO TRADE-FAS.

[7] Out of 338 observations Bermuda appears 146 times as the exporter (home), Cyprus appears 60 times and Antilles appears 59 times.

countries. Define as Y_j the income in country j (equal to expenditure by country j). Then, the consumer's utility maximization problem allows us to express the demand for each good as:

$$x_j(l) = \frac{\check{p}_j(l)^{-\epsilon}}{P_j^{1-\epsilon}} Y_j \tag{2.2}$$

where $\check{p}_j$ is the price of product l in country j and P_j is the standard CES ideal price index.[8]

As for technology, in country j the unit production cost of the firms is represented by a cost minimizing combination of inputs that costs $c_j a$, where c_j is country specific, while a is a firm-specific inverse indicator of productivity. Firms draw a randomly from a distribution $G(a)$, which is common across countries. The support for a is exogenously defined to be $[a_L, a_H]$.[9] There are no fixed production costs, hence firms never exit from the domestic market.

The market structure is characterized by usual monopolistic competition, hence the firms' profit maximization problem gives the optimal pricing rule as a constant mark-up over marginal cost:

$$p_{jj}(l) = \frac{c_j a}{\alpha} \tag{2.3}$$

where p_{jj} is the mill price of a variety produced in country j and sold in country j. There is no entry and the number of firms in country j is N_j.[10]

[8] Expressed by $P_j^{1-\epsilon} = \int_{l \in \Omega_j} \check{p}_j(l)^{1-\epsilon} dl$

[9] Also common across countries

[10] Like in HMR, but differently from HMY and Melitz(2003).

A domestic firm, besides serving the domestic market, can decide to serve foreign market i in two ways. If it decides to export, it has to bear a fixed cost $c_j f_{ij}^x$ and it is subject to an iceberg melting cost $\tau_{ij} > 1$.[11] The price in i of a good shipped from j will be therefore:

$$p_{ij}(l) = \tau_{ij} \frac{c_j a}{\alpha} \tag{2.4}$$

On the other hand, if the firm in country j decides to invest abroad, it has to bear a fixed cost $c_j f_{ij}^I$ but it does not have to pay the transport cost. Departing here from HMY, I assume that multinational operations involve higher costs than domestic operations due to a *cost disadvantage* for foreign firms in the purchase of intermediate inputs from local producers. Examples of how this cost disadvantage could arise include having less information about local markets or less experience dealing with local bureaucracy. Due to the presence of this cost disadvantage, p_{ii}^*, the price charged in country i by a multinational firm whose headquarters is located in country j will be:

$$p_{ii}^*(l) = \tau_{ij}^I \frac{c_i a}{\alpha} \tag{2.5}$$

τ_{ij}^I is the cost disadvantage over local producers, which is assumed to be increasing in the *cultural distance* between the two countries. The assumption of the presence of this cost disadvantage is a partial answer to the observation that "standard models of the proximity-concentration trade-off are missing an ingredient that would explain why the unit cost of serving foreign markets appears to rise in distance"(Yeaple, 2009).[12] τ_{ij}^I is defined

[11] As usual, $\tau_{jj} = 1$

[12] An alternative interpretation relies on the presence of monitoring costs, which increase the variable costs for affiliate of foreign companies over the marginal cost of domestic firms (Aizenman & Spiegel, 2007). In this case it is natural for cost to increase with distance.

for convenience to be a fraction of the transportation cost: $\tau_{ij}^I = \tau_{ij}^b$ with $b < 1$. The firms still face the concentration-proximity trade-off empirically documented in previous literature (Brainard, 1997).

Substituting the demand expression and the pricing rule into the expression for firms' profits and assuming a symmetric equilibrium, it is possible to express the *additional* profit that a firm gets from exporting as:

$$\pi_{ij}^x = (1-\alpha)\left(\frac{\tau_{ij} c_j a}{\alpha P_i}\right)^{1-\epsilon} Y_i - c_j f_{ij}^x \tag{2.6}$$

Notice the dependence of profits on firm-specific productivity a. Similarly, the additional operational profits for a firm that invests abroad can be expressed as

$$\pi_{ij}^I = (1-\alpha)\left(\frac{\tau_{ij}^b c_i a}{\alpha P_i}\right)^{1-\epsilon} Y_i - c_j f_{ij}^I \tag{2.7}$$

Following HMY, and calling $A_i = (1-\alpha)\frac{1}{(\alpha P_i)^{1-\epsilon}} Y_i$, I can re-write the previous expressions as:

$$\pi_{ij}^x = A_i(\tau_{ij} c_j)^{1-\epsilon} a^{1-\epsilon} - c_j f_{ij}^x \tag{2.8}$$

and

$$\pi_{ij}^I = A_i(\tau_{ij}^b c_i)^{1-\epsilon} a^{1-\epsilon} - c_j f_{ij}^I \tag{2.9}$$

Note that, with $\epsilon > 1$, the previous expressions are linear functions of a variable increasing in productivity. Figure 2.1 shows on the same graph equations (2.8) and (2.9), where I further impose two parameter restrictions:

$$(\tau_{ij} c_j)^{1-\epsilon} < (\tau_{ij}^b c_i)^{1-\epsilon} \quad (2.10)$$

$$\left(\frac{c_j}{c_i}\right)^{1-\epsilon} f_{ij}^I > \tau^{(1-b)(\epsilon-1)} f_{ij}^x \quad (2.11)$$

Eq (2.10) is needed to guarantee that we will observe FDI for some country-pairs. Equation (2.11) implies that FDI flows are observed only in the presence of trade flows, consistent with the evidence presented in section two.[13] It is clear from Figure 2.1 that there will be a productivity cut-off $(a_{ij}^x)^{1-\epsilon}$ below which the firm will not find it profitable to export. Most interestingly, though, there will be a second cut-off productivity $(a_{ij}^I)^{1-\epsilon}$, above which firms will prefer to invest abroad.

The two cut-offs are implicitly defined by the following conditions:

$$(1-\alpha)\left(\frac{\tau_{ij} c_j a_{ij}^x}{\alpha P_i}\right)^{1-\epsilon} Y_i = c_j f_{ij}^x \quad (2.12)$$

and

$$(1-\alpha)\frac{Y_i}{(\alpha P_i)^{1-\epsilon}}\left[\left(\tau_{ij}^b c_i\right)^{1-\epsilon} - (\tau_{ij} c_j)^{1-\epsilon}\right]\left(a_{ij}^I\right)^{1-\epsilon} = c_j\left(f_{ij}^I - f_{ij}^x\right) \quad (2.13)$$

In equation (2.12) the cutoff a_{ij}^x is defined as the productivity of the firm which is just indifferent between exporting or not, given that its additional profits from exporting are just enough to pay for the fixed costs. Equation (2.12), instead, defines the second cut-off

[13]Equation (11) is similar to the parameter restriction imposed in HMY.

a_{ij}^{I} as the productivity of the firm that is indifferent between serving the foreign market by exporting or by FDI. The reason for this indifference is that the additional profits are the same in the two cases.

The pattern of possibilities that emerges from the interaction between the two cut-offs implicitly identified by (2.12) and (2.13) and the exogenous support for the productivity draws is very rich and extends the possibilities allowed for by HMR. Figure 2.2 helps visualize the three possibilities. If $a_L^{1-\epsilon}$ [14] is lower than the trade productivity cut off, neither trade nor FDI flows will be observed between the two countries. If $a_L^{1-\epsilon}$ is between the two cut-offs, a fraction T_{ij} of firms will find it profitable to export, hence we will observe trade, but no FDI between the two countries. Finally, if $a_L^{1-\epsilon}$ is bigger than both cut-offs, we will observe both firms investing abroad (a fraction F_{ij} of them) and exporting (a fraction T_{ij}). In this case we will observe both FDI and bilateral trade flows. The three possible outcomes, hence, are fully consistent with the empirical evidence presented in section 2.2.

Finally, it is possible to derive expressions for the bilateral trade and investment flows. First, define two variables that represent the fraction of firms exporting and investing from country j to country i respectively:

$$T_{ij} = \begin{cases} \int_{a_{ij}^{I}}^{a_{ij}^{x}} a^{1-\epsilon} dG(a) & \text{if } a_L < a_{ij}^{I} \\ \int_{a_L}^{a_{ij}^{x}} a^{1-\epsilon} dG(a) & \text{if } a_{ji}^{I} < a_L < a_{ij}^{x} \\ 0 & \text{otherwise} \end{cases}$$

$$F_{ij} = \begin{cases} \int_{a_L}^{a_{ij}^{I}} a^{1-\epsilon} dG(a) & \text{if } a_L < a_{ij}^{I} \\ 0 & \text{otherwise} \end{cases}$$

[14] The level of productivity of the most productive firm.

Then, the *value of imports* in country i from country j is given by:

$$M_{ij} = \left(\tau_{ij}\frac{c_j}{\alpha P_i}\right)^{1-\epsilon} Y_i N_j T_{ij} \tag{2.14}$$

and the *value of the foreign affiliate sales (FAS)* in country i from country j would be given by:

$$FAS_{ij} = \left(\tau_{ij}^{b}\frac{c_i}{\alpha P_i}\right)^{1-\epsilon} Y_i N_j F_{ij} \tag{2.15}$$

It is important to stress that equation (2.15) refers to FAS more than to FDI. This is the reason why in the implementation of the empirical methodology presented in the next section I will only use the dataset on bilateral FAS coming from the ORBIS firm-level dataset.

2.4 Empirical framework

Productivity is assumed to be drawn from a Pareto distribution, hence $G(a) = \frac{a^k - a_L^k}{a_H^k - a_L^k}$. Analogously to HMR, F_{ij} can be found as:

$$F_{ij} = \frac{k a_L^{k-\epsilon+1}}{(k-\epsilon+1)(a_H^k - a_L^k)} W_{ij}^1 \tag{2.16}$$

where

$$W_{ij}^1 = max\left[\left(\frac{a_{ji}^I}{a_L}\right)^{k-\epsilon+1} - 1, 0\right] \tag{2.17}$$

Things are more complicated, instead, for the trade equation, since now the fraction of exporting firms depends on whether there are firms from country j investing in country i or not. In particular, we would have

$$T_{ij} = \begin{cases} \frac{ka_L^{k-\epsilon+1}}{(k-\epsilon+1)(a_H^k - a_L^k)} W_{ij}^2 & \text{if } F_{ij} = 0 \\ \frac{ka_L^{k-\epsilon+1}}{(k-\epsilon+1)(a_H^k - a_L^k)} W_{ij}^3 & \text{if } F_{ij} \neq 0 \end{cases}$$

where

$$W_{ij}^2 = max \left[\left(\frac{a_{ij}^x}{a_L} \right)^{k-\epsilon+1} - 1, 0 \right] \tag{2.18}$$

and

$$W_{ij}^3 = \left[\left(\frac{a_{ij}^x}{a_L} \right)^{k-\epsilon+1} - \left(\frac{a_{ij}^I}{a_L} \right)^{k-\epsilon+1} \right] \tag{2.19}$$

From (2.15), it is possible to express the foreign affiliate sales equation in its log-linear form as:

$$fas_{ij} = (\epsilon - 1)ln\alpha - (\epsilon - 1)lnc_i + n_j + (\epsilon - 1)p_i + y_i + b(1 - \epsilon)ln\tau_{ij} + f_{ij} \tag{2.20}$$

where the lower case variables represent the natural logarithm of the corresponding upper case variables. Following HMR, I assume the following functional form for τ_{ij}:

$$\tau_{ij}^{\epsilon-1} = D_{ij}^{\gamma} e^{-u_{ij}^1} \tag{2.21}$$

where D_{ij} is an indicator of the economic distance between j and i and u_{ij}^1 is assumed to be i.i.d. normally distributed with mean zero and variance σ_1^2. Given (21) it is possible to derive the following estimation equation (2.22):

$$fas_{ij} = \theta_0 + \Psi_j^I + \Upsilon_i^I - \gamma_1 d_{ij} + w_{ij}^1 + bu_{ij}^1 + \nu_{ij} \tag{2.22}$$

where $\Psi_j^I = n_j$ is a home country fixed effect and $\Upsilon_i^I = -(\epsilon-1)lnc_i + (\epsilon-1)p_i + y_i$ is a host country fixed effect, $\gamma_1 = b\gamma$ and θ_0 contain the elements in F_{ij} besides W_{ij}^1. bu_{ij}^1 is i.i.d. normally distributed with mean zero and variance $b^2\sigma^2$. ν_{ij} represents measurement error in the dependent variable. It is assumed to be i.i.d. distributed with mean zero and variance σ_ν^2.

On the other hand, taking logs of equation (2.14) and taking into account equations (2.18), (2.19) and (2.21), it is possible to express the trade flow equation as the following estimable equation:

$$m_{ij} = \theta_1 + \Psi_j^x + \Upsilon_i^x - \gamma d_{ij} + w_{ij}^s + u_{ij}^1 \tag{2.23}$$

where $\Psi_j^x = n_j - (\epsilon-1)lnc_j$ is an exporter fixed effect, $\Upsilon_i^x = (\epsilon-1)p_i + y_i$ is an importer fixed effect and θ_1 includes all the elements in T_{ij} besides W^s, with $s = [2,3]$.

Looking at equations (2.22) and (2.23), three things are worth noticing. First, not taking into account the term W^s might lead to inconsistent estimates of all the coefficients. Second, the model has clear predictions regarding the relative magnitude of the distance coefficients in the trade and FAS equations: they are expected to be higher in the trade flows equation. Third, the form of the estimating equation for the trade flows will differ according to whether FAS are observed or not, since in the presence (or absence) of FAS changes the correction term for firm heterogeneity (w_{ij}). Not recognizing the option for a firm to serve a foreign market by directly investing instead of exporting leads to an overestimate of the fraction of exporting firms that could affect the estimates of all the coefficients in equation (2.23). The importance of this possible bias is ultimately an empirical question.

The next subsection outlines a two-stage procedure aimed at consistently estimating equations (2.22) and (2.23).

2.4.1 First Stage: Selection

As explained before, this framework allows for endogenous selection into Export and FAS). The first stage evaluates the self-selection problem and can be best understood as a three-steps process.

The first step consists of defining adequate latent variables. In particular, analogously to HMR, I can define a latent variable Z_{ij}^x which determines whether we should observe trade flows from country j to country i as follows:

$$Z_{ij}^x = \frac{(1-\alpha)\left(\frac{\tau_{ij}c_j}{\alpha P_i}\right)^{1-\epsilon} Y_i a_L^{1-\epsilon}}{c_j f_{ij}} \tag{2.24}$$

Z_{ij}^x represents the ratio of the variable export profit for the most productive firms to the fixed export costs where $f_{ij}^x = f_{ij}$. Clearly, we would observe export from j to i only if $Z_{ij}^x > 1$.

For simplicity, I assume the investment fixed cost to be a multiple of the trade fixed cost, i.e. $f_{ij}^I = q f_{ij}$ with $q > 1$. Then, starting from equation (2.12), it is possible to define a second latent variable Z_{ij}^I, which is the ratio of the difference in the variable profits from investment and export to the difference in the fixed costs:

$$Z^I_{ij} = \frac{(1-\alpha)\frac{Y_i}{(\alpha P_i)^{1-\epsilon}}\left[\left(\tau^b_{ij}c_i\right)^{1-\epsilon} - (\tau_{ij}c_j)^{1-\epsilon}\right](a_L)^{1-\epsilon}}{(q-1)c_j f_{ij}} \tag{2.25}$$

If $Z^I_{ij} > 1$ we should observe both trade and FAS between countries. Now it is convenient to define a third *auxiliary* latent variable Z_{ij}, representing the ratio of the variable profits from investment to the fixed cost of investment for the most productive firm:

$$Z_{ij} = \frac{(1-\alpha)\left(\frac{\tau^b_{ij}c_i}{\alpha P_i}\right)^{1-\epsilon} Y_i a_L^{1-\epsilon}}{c_j q f_{ij}} \tag{2.26}$$

In other words, $Z_{ij} > 1$ implies that the most productive firm *could* profitably invest abroad, even though it *might* prefer to export instead if its productivity is lower than $\left(a^I_{ij}\right)^{1-\epsilon}$. Equation (2.26) is particularly helpful because it allows me to express the other two latent variables as functions of Z. In fact, from (2.24) and (2.26) we can see how:

$$Z^x_{ij} = Z_{ij} q \left(\frac{\tau_{ij}c_j}{\tau^b_{ij}c_i}\right)^{1-\epsilon} \tag{2.27}$$

Hence, we would observe trade between country i and j if $Z_{ij} > \Delta_1$, where $\Delta_1 = \frac{1}{q}\left(\frac{\tau_{ij}c_j}{\tau^b_{ij}c_i}\right)^{\epsilon-1}$, which according to equation (11) is a quantity smaller than one. Importantly, I'm assuming here that $\left(\frac{\tau_{ij}c_j}{\tau^b_{ij}c_i}\right)^{\epsilon-1}$ is a constant (smaller than 1 by equation (2.10)). The economic meaning of this assumption is that the variable cost of a firm with productivity a who decides to invest abroad is a fraction of the variable cost that the same firm faces if it decides to export abroad instead.[15]

[15]Essentially I'm parameterizing the well accepted existence of a proximity-concentration trade-off by making the fixed cost of investing a multiple of the fixed cost of exporting and the variable cost of investing

In a similar fashion, from equations (2.24) (2.25) (2.26) and (2.27) we can derive

$$Z_{ij}^{I}=\frac{q}{q-1}Z_{ij}-\frac{1}{q-1}Z_{ij}^{x}=\frac{q\Delta_1-1}{\Delta_1\left(q-1\right)}Z_{ij} \tag{2.28}$$

Hence we will observe FAS between country j and country i if $Z_{ij}^{I}>1$, or $Z_{ij}>\Delta_2$ where $\Delta_2=\frac{\Delta_1(q-1)}{q\Delta_1-1}$, which given our parameter restrictions is a quantity bigger than 1. In order to derive an estimable equation from (2.26), I assume that fixed trade costs are stochastic due to unmeasured frictions. Specifically, I assume that:

$$f_{ij}=e^{\kappa\phi_{ij}-u_{ij}^{2}} \tag{2.29}$$

where ϕ_{ij} are a series of factors that influence the fixed costs of exporting (possibly common to the elements that enter in the definition of economic distance) and u_{ij}^{2} is assumed to be i.i.d. normally distributed with mean zero and variance σ_2^2. With this assumption, I can express equation (26) as

$$z_{ij}=\theta_2+\Psi_j+\Upsilon_i-\gamma d_{ij}-\kappa\phi_{ij}+e_{ij} \tag{2.30}$$

where Ψ_j are exporter/home fixed effects, Υ_i are importer/host fixed effects and $e_{ij}=bu_{ij}^{1}+u_{ij}^{2}$ is i.i.d. normally distributed with mean zero and variance $\sigma_{e_{ij}}^{2}=b^2\sigma_1^2+\sigma_2^2$. Notice also that, given the definitions of Δ_1 and Δ_2, it is possible to express the latent variables $z_{ij}^{x}=lnZ_{ij}^{x}=z_{ij}-\delta_1$ and $z_{ij}^{I}=lnZ_{ij}^{I}=z_{ij}-\delta_2$, where $\delta_1=ln\Delta_1$ and $\delta_2=ln\Delta_2$.

The dependence of both Z_{ij}^{I} and Z_{ij}^{x} on Z_{ij} allows me to use an *ordered Probit* model to control for selection and heterogeneity. The second step in the procedure is to define an

a fraction of the variable cost of exporting.

ordered outcome variable $GLOBAL_{ij}$, which can take values zero ($TRADE_{ij} = 0, FDI_{ij} = 0$), one ($TRADE_{ij} = 1, FDI_{ij} = 0$) or two ($TRADE_{ij} = 1, FDI_{ij} = 1$), consistent with the pattern showed in section two.

Following HMR, I do not impose unitary variance on the error process and I divide equation (30) by $\sigma_{e_{ij}}$. It is thus possible to obtain the following ordered Probit model:

$$z_{ij}^* = \theta_2^* + \Psi_j^* + \Upsilon_i^* - \gamma^* d_{ij} - \kappa^* \phi_{ij} + e_{ij}^* \tag{2.31}$$

with

$$\begin{cases} TRADE_{ij} = 0, FDI_{ij} = 0 & \text{if } z_{ij}^* < \delta_1^* \\ TRADE_{ij} = 1, FDI_{ij} = 0 & \text{if } \delta_1^* < z_{ij}^* < \delta_2^* \\ TRADE_{ij} = 1, FDI_{ij} = 1 & \text{if } z_{ij}^* > \delta_2^* \end{cases}$$

where the starred variables and parameters represent the original variables and parameters divided by the relevant standard deviation and e_{ij}^* is now i.i.d. unit normally distributed. Importantly, as stressed by HMR, the selection equation is derived from firm-level decision and does not contain the unobserved terms W_{ij}^s.

Finally, as a third step, from the ordered Probit estimates it is possible to recover consistent estimates of W_{ij}^s,$s = [1, 2, 3]$, which can then be used in the flow equation to correct for heterogeneity. Let $\hat{p}_{ij}^0$ be the predicted probability of not observing trade nor FAS flows between countries j and i. Then, $\hat{z}_{ij}^{x*} = -\Phi^{-1}\left(\hat{p}_{ij}^0\right)$ is the predicted value of the latent variable $z_{ij}^{x*} = \frac{z_{ij}^x}{\sigma_{e_{ij}}}$.[16] In a similar fashion, calling $\hat{p}_{ij}^2$ the predicted probability of

[16]To see this, define for simplicity $\theta_2^* + \Psi_j^* + \Upsilon_i^* - \gamma^* d_{ij} - \kappa^* \phi_{ij} = x_{ij}\beta^*$. Then $p_{ij}^0 = Prob\left[x_{ij}\beta^* + e_{ij}^* < \delta_1^*\right] = \Phi\left(\delta_1^* - x_{ij}\beta^*\right)$. Hence $-\Phi^{-1}\left(\hat{p}_{ij}^0\right) = (x_{ij}\hat{\beta^*} - \delta_1^*) = \hat{z}_{ij}^{x*}$.

observing both trade and FAS between country i and j, $\hat{z}_{ij}^{I*} = \Phi^{-1}\left(\hat{p}_{ij}^2\right)$ is the predicted value of the latent variable $z_{ij}^{I*} = \frac{z_{ij}^I}{\sigma_{e_{ij}}}$.[17] With these two predicted values, we can obtain consistent estimates of W_{ij}^s, $s = [1, 2, 3]$ as follows:

$$W_{ij}^1 = max\left[\left(Z_{ij}^{I*}\right)^{\zeta} - 1, 0\right] \tag{2.32}$$

$$W_{ij}^2 = max\left[\left(Z_{ij}^{x*}\right)^{\zeta} - 1, 0\right] \tag{2.33}$$

$$W_{ij}^3 = \left[\left(Z_{ij}^{x*}\right)^{\zeta} - \left(Z_{ij}^{I*}\right)^{\zeta}\right] \tag{2.34}$$

with $\zeta = \sigma_{e_{ij}} \frac{k-\epsilon+1}{\epsilon-1}$.[18]

2.4.2 Second Stage: FAS and TRADE Log-Linear Equations

In order to consistently estimate equations (2.22) and (2.23), I need to correct for both heterogeneity and selection. This requires the estimation of the different expected values for w_{ij} for the cases of only trade or both trade and FAS flows between countries; hence, I need $E\left[w_{ij}^1|., GLOBAL_{ij} = 2\right]$, $E\left[w_{ij}^2|., GLOBAL_{ij} = 1\right]$ and $E\left[w_{ij}^3|., GLOBAL_{ij} = 2\right]$. Moreover, I need to evaluate the expected values of the error terms in the different cases, that is to say I also need $E\left[u_{ij}^1|., GLOBAL_{ij} = 1\right]$ and $E\left[u_{ij}^1|., GLOBAL_{ij} = 2\right]$. Analogously to HMR, I exploit here the dependence of all these terms on e_{ij}^*, which is unit normal. In particular, using the properties of the truncated standard normal, I can derive:

[17] Defining $x_{ij}\beta^*$ as in in the previous note, $p_{ij}^2 = Prob\left[x_{ij}\beta^* + e_{ij}^* > \delta_2^*\right] = \Phi\left(x_{ij}\beta^* - \delta_2^*\right)$. Hence $\Phi^{-1}\left(\hat{p}_{ij}^2\right) = (x_{ij}\hat{\beta^*} - \delta_2^*) = \hat{z}_{ij}^{I*}$.

[18] See Equations (2.13) and (2.25) to derive equation (2.32) and equations (2.11) and (2.23) to get equations (2.33) and (2.34).

$$
\begin{aligned}
E\left[e_{ij}^{*}|.,z_{ij}^{*}>\delta_2^{*}\right] &= \frac{\phi\left(\hat{z}_{ij}^{I*}\right)}{\Phi\left(\hat{z}_{ij}^{I*}\right)}=\hat{\eta}_{ij}^{1} &(2.35)\\
E\left[e_{ij}^{*}|.,z_{ij}^{*}>\delta_1^{*}\right] &= \frac{\phi\left(\hat{z}_{ij}^{x*}\right)}{\Phi\left(\hat{z}_{ij}^{x*}\right)}=\hat{\eta}_{ij}^{2} &(2.36)\\
E\left[e_{ij}^{*}|.,\delta_1^{*}<z_{ij}^{*}<\delta_2^{*}\right] &= \frac{\phi\left(-\hat{z}_{ij}^{x*}\right)-\phi\left(-\hat{z}_{ij}^{I*}\right)}{\Phi\left(\hat{z}_{ij}^{I*}\right)-\Phi\left(\hat{z}_{ij}^{x*}\right)}=\hat{\eta}_{ij}^{3} &(2.37)
\end{aligned}
$$

where $\phi()$ and $\Phi()$ are the p.d.f. and the c.d.f. of the standard normal. Using equation (2.35), (2.36) and (2.37) I can get consistent estimates for w_{ij} as follows:

$$
\begin{aligned}
\hat{w}_{ij}^{1} &= ln\left\{exp\left[\zeta\left(\hat{z}_{ij}^{I*}+\hat{\eta}_{ij}^{1}\right)\right]-1\right\} &(2.38)\\
\hat{w}_{ij}^{2} &= ln\left\{exp\left[\zeta\left(\hat{z}_{ij}^{x*}+\hat{\eta}_{ij}^{2}\right)\right]-1\right\} &(2.39)\\
\hat{w}_{ij}^{3} &= ln\left\{exp\left[\zeta\left(\hat{z}_{ij}^{x*}+\hat{\eta}_{ij}^{3}\right)\right]-exp\left[\zeta\left(\hat{z}_{ij}^{I*}+\hat{\eta}_{ij}^{1}\right)\right]\right\} &(2.40)
\end{aligned}
$$

Hence, it is possible to consistently estimate equation (23) using the following transformation:

$$
fas_{ij}=\theta_0+\Psi_j^{I}+\Upsilon_i^{I}-\gamma_1 d_{ij}+ln\left\{exp\left[\zeta\left(\hat{z}_{ij}^{I*}+\hat{\eta}_{ij}^{1}\right)\right]-1\right\}+\beta^{1}\hat{\eta}_{ij}^{1}+e_{ij}^{1} \qquad (2.41)
$$

where e_{ij}^{1} is an i.i.d error for which $E\left[e_{ij}^{1}|.,GLOBAL_{ij}=2\right]=0$. Equation (2.41) can be estimated via non-linear least squares (as in HMR) or through Maximum Likelihood (as I will do in the next section).

Consistent estimation of equation (2.22) now depends on whether we also observe investment flows between the two countries. If only trade is observed, then it is possible to estimate the trade flows gravity-type equation as:

$$m_{ij} = \theta_1 + \Psi_j^x + \Upsilon_i^x - \gamma d_{ij} + ln\left\{exp\left[\zeta\left(\hat{z}_{ij}^{x*} + \hat{\eta}_{ij}^2\right)\right] - 1\right\} + \beta^2\hat{\eta}_{ij}^2 + e_{ij}^2 \tag{2.42}$$

where e_{ij}^2 is an i.i.d error for which $E\left[e_{ij}^2|., GLOBAL_{ij} = 1\right] = 0$. On the other hand, if FDI (FAS) are also observed between countries, then the correct way to estimate equation (2.22) becomes:

$$m_{ij} = \theta_1 + \Psi_j^x + \Upsilon_i^x - \gamma d_{ij} + ln\left\{exp\left[\zeta\left(\hat{z}_{ij}^{x*} + \hat{\eta}_{ij}^3\right)\right] - exp\left[\zeta\left(\hat{z}_{ij}^{I*} + \hat{\eta}_{ij}^1\right)\right]\right\} + \beta^3\hat{\eta}_{ij}^3 + e_{ij}^3 \tag{2.43}$$

where e_{ij}^3 is an i.i.d error for which $E\left[e_{ij}^3|., GLOBAL_{ij} = 2\right] = 0$ and is potentially correlated with e_{ij}^1. Importantly, using the correction terms contained in equation (42) to address cases where also FAS are present implies a possible omitted variable bias (given that the correct correction terms that should be applied are the one of equation (2.43)). The relevance of this potential problem is ultimately an empirical question.[19]

Before proceeding to the results, it is probably useful to briefly summarize the notationally-intensive procedure. Essentially, I am proposing a two-stage procedure for the estimation of bilateral trade and FAS flows. In the first stage, the definition of convenient latent variables allows me to describe the self-selection of heterogeneous firms into trade and FAS

[19]The quantitative magnitude of this potential problem is found to be small in the dataset used here.

with an ordered Probit model. From the ordered Probit estimates, it is possible to back out variables that allow me to correct the flows equations for selection and for the fraction of exporting/investing firms.

An important caveat to this methodology is the same as that which has been noted about the original HMR methodology: the possible inconsistency from using in the correction terms elements correlated to the errors. As pointed out by Santos Silva and Tenreyro (2008), though, the correction method proposed can be considered approximately right for many practical situations.

2.5 Results

This section reports the results obtained from applying the methodology introduced in the previous section to the dataset on bilateral trade and foreign affiliate sales in manufacturing introduced in section 2.2.

As regressors, I use variables from the distance dataset from the CEPII. This dataset provides me with data for a large sample of countries and includes variables such as geographic distance and dummies for common border, presence of a regional trade agreement common colonial past and common language.[20] I'll first present the OLS estimates and then move on to those from the two-stage procedure. All the trade, FDI and FAS flows values are measured in 2000 US dollars, converted using the US CPI.

I also add time fixed effects in all the equations. While the model presented in section three is static and should ideally be tested only on a large cross section of data, by pooling the data from several years and including year fixed effects I increase the number of

[20]The data on RTA are taken from the dataset by Head et al (2010).

observations available.

2.5.1 OLS Estimates

Table 2.3 contains the results obtained using OLS techniques on the bilateral FAS and trade in the manufacturing sector for the period 2000-2006.The first column includes the results obtained for the FAS equation and the second column includes the results obtained for the trade equation using only the observations where FAS were not observed. Finally, the third column reports the results obtained for the trade equation including only those observations where also FAS were observed.

Two interesting results emerge: First, the coefficient on distance is smaller in the FAS equation than in the trade equations. Second, the coefficient on distance in the trade (no FAS) equation is higher than the coefficient on distance in the trade (with FAS) equation. The first result is consistent with the proposed assumption of the theoretical model in section three. The difference in the two coefficients can be interpreted as a cost disadvantage faced by foreign firms that is less important than the transport cost needed to ship goods internationally. The second result, which is clearly not modeled here, could potentially be explained by the fact that countries also experiencing FDI are generally richer and more integrated. The *composition of trade* between those countries, hence, might be inherently characterized by goods less sensitive to distance.

Consistently with previous evidence, the presence of a common border, a common colonial origin, and an RTA appear to have positive effects on both the FAS and the trade bilateral flows. Sharing a common language appears to have a positive impact on bilateral trade flows but not on bilateral FAS flows.

2.5.2 Two-Stage Estimation

Table 2.4 reports the results of the ordered Probit regression. Given that all the coefficients have been divided by σ_e, the quantitative magnitude is not very revealing. Distance, as expected, decreases the probability of observing bilateral trade and FAS flows between countries. Colonial links and common language seem to be significant variables in determining the probability of observing positive trade/investment flows. The presence of a common border displays a positive coefficient, which is not statistically significant.[21]

Moreover, in order to avoid relying on identification through functional form in the estimation of equations (2.41), (2.42) and (2.43), it is necessary to include in the first stage a variable that is excluded in the second stage. Following HMR, I use a common religion variable as the excluded variable.[22] As Table 2.4 shows, common religion is a significant factor in determining the probability of observing trade and investment flows between countries, thus making it a useful excluded variable.[23]

Table 2.5 reports the results obtained for the bilateral FAS flow equation (2.41). The first columns report the OLS estimates to ease the comparison (equivalent to the first columns of Table 2.3). The second column of results are obtained by estimating equation (2.41) through maximum likelihood. The coefficient on distance is less than half of its OLS counterpart in the specification that corrects for heterogeneity and selection. Thus the result found by HMR for trade is found to be valid also for bilateral foreign affiliate

[21]this last result is reminiscent of HMR, who find a negative coefficient on border in their first stage regression. HMR proposed as justification the effects of territorial border conflicts that suppresses trade between neighbors.

[22]Expressed as the probability that two randomly picked individuals in the two countries in 1996 belong to the same religion.

[23]Although not a perfect one. I plan in future work to explore the increased availability of possible excluded variables suitable for this framework.

sales. The OLS coefficient does not properly distinguish between the impact of distance on the extensive margin of the international activities (the number of firms able to invest, in this case) and the amount of their sales in the foreign market (the intensive margin). The importance of correcting the estimates for the presence of firm heterogeneity is witnessed by the coefficients on the corrections term ζ and $\hat{\eta}^1$, which are both highly statistically significant. The coefficient on colonial origin is half of what obtained with OLS (and become statistically insignificant), while the coefficient on RTA is reduced by roughly one third. The coefficient on border is also reduced, though to a lessen degree.

Table 2.6 reports the results obtained for the trade equations in the absence of FAS. The OLS estimates are reported for comparison. Also in the case of trade, the coefficient obtained with the two-stage estimation procedure are smaller than the one obtained with OLS. The drop of the coefficients on distance and common border are between 10% and 15%, while the drops of the coefficients on RTA and common colonial origin are more substantial (of the order of 30%). Overall, the results in table 2.6 are broadly consistent with what previously found by HMR. The importance of correcting the estimates for selection and heterogeneity is confirmed by the strong statistical significance of both correction terms.

Finally, Table 2.7 reports the results of the trade equation obtained considering the cases where also FAS were present. The first column report the OLS estimates for comparison. The second column contain the results obtained by applying the correction terms included in equation (2.42), which do not take into account of the presence of FAS while the third column contains the correction terms that take into account of FAS.[24] The coefficients on

[24]The coefficients obtained in column two are not exactly the coefficients that would be derived by an HMR-type procedure because the correction terms still come from the first stage of the procedure proposed in this paper (the order probit). Conceptually, it would be necessary to compare the coefficients of column 3 with what would be obtained using a probit model for trade flows as first stage. Unfortunately, the small

all the variables are lower than the OLS estimates in both cases. However, the coefficients obtained taking into account of the FAS tend to be higher than the ones obtained without taking into account of the FAS. The intuition for this result is that not taking into account the the FAS implies overestimating the fraction of firms that export (including also some firms that actually invest and serve the foreign market through sales by its affiliate). Overestimating the fraction of exporting firms imply underestimating the importance of factors such as distance, RTA and so on on the intensive margin. From a quantitative point of view, however, the difference in the coefficients is not very large. Overall, I conclude that failing to properly take into account the presence of FAS when correcting the aggregate trade flow equation for selection and heterogeneity has a modest impact on the estimated coefficients.

2.6 Conclusion

I have uncovered a new pattern in the data, namely that bilateral FDI and FAS are almost never observed in the absence of bilateral trade flows. I developed a model with implications for aggregate trade flows and foreign affiliate sales that are consistent with this pattern. I proposed a two-stage methodology, structurally derived from the model, to estimate the trade and the FAS equations consistently. The main results of the analysis are as follows: 1) The impact of distance, border and regional trade agreements on bilateral foreign affiliate sales becomes substantially smaller after controlling for selection and firms' heterogeneity (hence separating the impact on the extensive versus the intensive margin).

number of zeroes in the sample makes impossible to estimate a first stage fully consistent with the HMR procedure. This is a caveat that must be considered when examining the results.

2) The same "attenuation result" is found also for the trade equations, consistently with HMR. 3) When FAS are observed, failing to take this into account when correcting for heterogeneity and selection in the trade equations leads to differences in the estimated coefficients.

Interesting directions for related future research include using the methodology proposed here on more disaggregated data (say at the level of single industries) to uncover possible differences across different sectors.

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Table 2.1: **Selection in Aggregate FDI and TRADE, 2000-2007**

	No Trade	**Trade**	**Total**
No FDI	2,671	14,978	17,649
FDI	**351**	9,884	10,235
Total	3,022	24,862	27,884

Table 2.2: **Selection in FAS and TRADE, Manufacturing, 2000-2006**

	No Trade	**Trade**	**Total**
No FAS	407	4,749	5,126
FAS	**338**	4,082	4,420
Total	745	8,831	9,576

Table 2.3: **OLS results, FAS and Trade, Manufacturing, 2000-2006**

Dep Variable	**FAS**	**trade NO FAS**	**trade FAS**
distance	-0.325***	-0.935***	-0.739***
	(0.059)	(0.025)	(0.022)
border	0.789***	0.303***	0.328***
	(0.135)	(0.070)	(0.049)
rta	0.644***	0.535***	0.718***
	(0.135)	(0.047)	(0.049)
colonial	0.442**	0.739***	0.401***
	(0.137)	(0.065)	(0.050)
language	-0.038	0.412***	0.328***
	(0.113)	(0.045)	(0.041)
Imp/Host FE	Yes	Yes	Yes
Exp/Home FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
R-squared	0.556	0.869	0.921
N	4082	4749	4082

Standard Errors in Parenthesis

*,**,*** Statistically Significant at 10%, 5% and 1%

Table 2.4: **Ordered Probit Results**

distance	-0.171***
	(0.037)
border	0.052
	(0.099)
rta	0.288***
	(0.074)
colonial	0.318***
	(0.094)
language	0.493***
	(0.068)
religion	**0.894*** **
	(0.110)
Imp/Host FE	Yes
Exp/Home FE	Yes
Year FE	Yes
pseudo R squared	0.57
N	9238

Standard Errors in Parenthesis

*,**,*** Statistically Significant at 10%, 5% and 1%

Table 2.5: **FAS Results, Manufacturing, 2000-2006**

Technique	**OLS**	**ML**
distance	-0.325***	-0.149**
	(0.059)	(0.074)
border	0.789***	0.683***
	(0.135)	(0.133)
rta	0.644***	0.473***
	(0.135)	(0.141)
colonial	0.442***	0.226
	(0.137)	(0.157)
language	-0.038	-0.486***
	(0.113)	(0.157)
ζ		1.252***
		(0.243)
$\hat{\eta}^1$		0.721***
		(0.278)
Imp/Host FE	Yes	Yes
Exp/Home FE	Yes	Yes
Year FE	Yes	Yes
R-squared	0.55	
N	4082	3968

Standard Errors in Parenthesis

*,**,*** Statistically Significant at 10%, 5% and 1%

Table 2.6: **Trade Results with No FAS, Manufacturing, 2000-2006**

Technique	**OLS**	**ML**
distance	-0.935***	-0.798***
	(0.025)	(0.030)
border	0.303***	0.274***
	(0.070)	(0.068)
rta	0.535***	0.379***
	(0.047)	(0.050)
colonial	0.739***	0.469***
	(0.065)	(0.072)
language	0.412***	0.067
	(0.045)	(0.064)
ζ		0.451***
		(0.107)
$\hat{\eta}^2$		-0.671***
		(0.227)
Imp/Host FE	Yes	Yes
Exp/Home FE	Yes	Yes
Year FE	Yes	Yes
R-squared	0.86	
N	4749	4749

Standard Errors in Parenthesis

*,**,*** Statistically Significant at 10%, 5% and 1%

Table 2.7: **Trade results with FAS, Manufacturing, 2000-2006**

Technique	**OLS**	**ML-HMR**	**ML**
distance	-0.739***	-0.571***	-0.584***
	(0.022)	(0.027)	(0.028)
border	0.328***	0.256***	0.266***
	(0.049)	(0.049)	(0.049)
rta	0.718***	0.529***	0.548***
	(0.049)	(0.051)	(0.052)
colonial	0.401***	0.080	0.096
	(0.050)	(0.057)	(0.059)
language	0.328***	-0.107*	-0.069
ζ		0.766***	0.724***
		(0.080)	(0.213)
$\hat{\eta}^2$		-11.43***	
		(1.36)	
$\hat{\eta}^3$			-0.764***
			(0.383)
Imp/Host	Yes	Yes	Yes
Exp/Home FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
R-squared	0.917		
N	4082	3971	3968

Standard Errors in Parenthesis

*,**,*** Statistically Significant at 10%, 5% and 1%

Figure 2.1: **Self-Selection into Export and FDI**

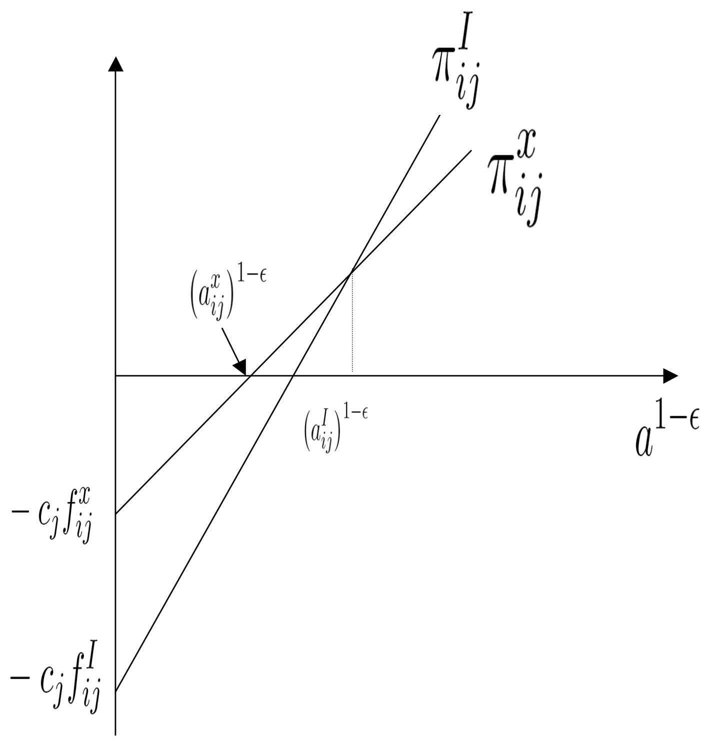
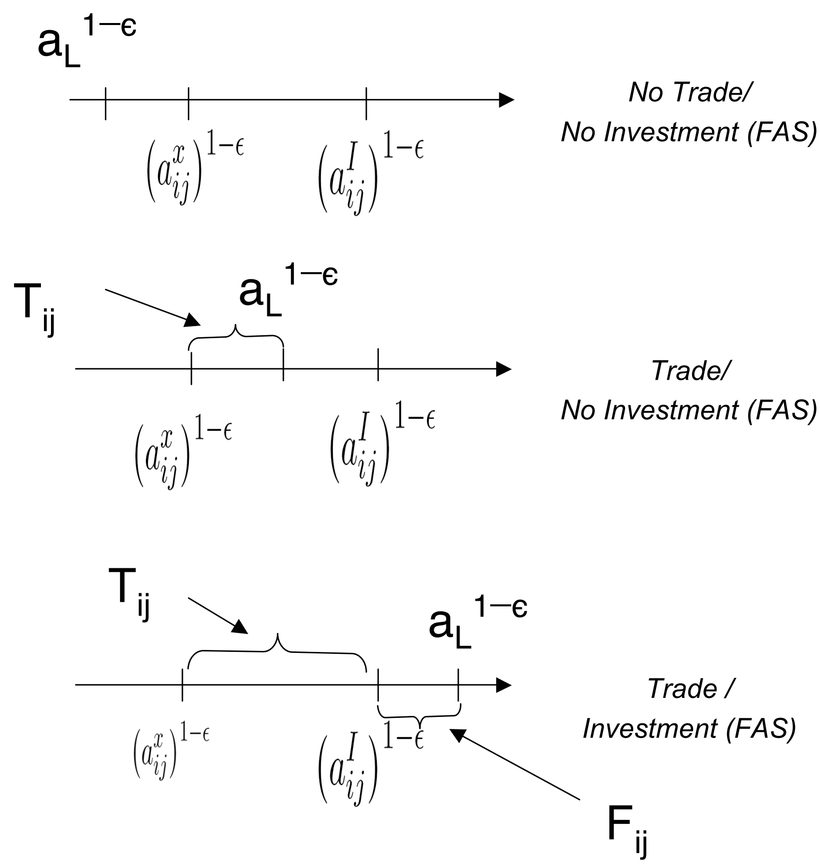


Figure 2.2: **Interaction between Bilateral Trade and FDI Thresholds and Most Productive Firm's Productivity: Three Possible Cases**



Chapter 3

Some Evidence on the Importance of the Sticky Wages

3.1 Introduction

It is difficult to explain the estimated real effects of monetary policy shocks without assuming that some nominal variables adjust sluggishly. In the *General Theory*, Keynes (1936) assumed that nominal wages were rigid, and thus that expansionary monetary policy would reduce real wages and increase employment and output. Fischer (1977) and Taylor (1980) showed that nominal wage contracts would have similar effects even in explicitly dynamic models with rational expectations. Recent macro-econometric models have typically followed the important contribution of Erceg, Henderson, and Levin (2000), and assumed that both prices and nominal wages are slow to adjust.

The large number of recent models with such features has inspired researchers to examine micro data on the frequency of price changes for individual products, with notable papers by Bils and Klenow (2004) and Nakamura and Steinsson (2008). However, to date there has been little research using micro data to estimate the rigidity of nominal wages — even though Christiano, Eichenbaum, and Evans (CEE 2005) find that nominal wage rigidity is more important than nominal price rigidity for explaining the dynamic effects of monetary policy shocks.

Our paper attempts to address this gap in the literature. The lack of previous work on the business cycle implications of nominal wage rigidity using micro data may be due in part to a lack of suitable datasets. We provide evidence about the frequency of wage adjustment in the United States using data from the Survey of Income and Program Participation (SIPP). The SIPP is a survey run by the Bureau of Labor Statistics (BLS). It provides individual wage histories for a large and representative sample that is followed for a period of 24 to 48 months. Importantly, the individuals are interviewed every four months. These data allow us to examine wage changes using high-frequency data. (Most previous work on nominal wage rigidity using U.S. micro data has used the PSID, which is an annual survey and thus less useful for high-frequency analysis. Other well-known sources of micro wage data, the CPS and the Employment Cost Index (ECI), do not provide sufficiently long time-series data on individual wages and thus cannot be used for our purpose.) We use the longest panel of the SIPP for which complete data are available: the 1996 panel (run from March 1996 to February 2000).

In our main results, we focus on the frequency of nominal wage adjustments disregarding employment history. This is arguably the concept that is most relevant for macro models with nominal wage rigidities, particularly medium-scale DSGE models *a la* CEE (2005). The reason is that most business-cycle models with nominal wage rigidity follow Blanchard and Kiyotaki (1987) and assume that all workers are monopolistically competitive suppliers of differentiated labor services. In this framework, the worker sets the wage, and revises it occasionally on his/her own schedule, thus making the sequence of wages the relevant series to examine, regardless of employment history.

We use as our baseline the results for hourly workers (or wage earners) who reported their hourly wages to the SIPP interviewer. The reason is that computing wages as hourly earnings increases measurement error. For the baseline results we chose to focus on the statistic measured with least error, the hourly wage, at the cost of making the sample less representative. However, we also present results for the sample of salaried workers, using their monthly earnings as their "wage" measure. By reporting the results both for hourly workers and for salaried workers, we leave the decision of the "right number" for macroeconomics to individual researchers who may be interested in calibrating their models using our estimates.

Regardless of the sample used, it is clear that the data are contaminated with a significant amount of measurement error. This is a disadvantage of working with data on individual wages, which in U.S. survey data are always self-reported.[1] We deal with this problem by applying to the reported wage and earnings series the correction for measurement error introduced by Gottschalk (2005), who built upon the work of Bai and Perron (1998 and 2003). The application uses the identifying assumption that wages are not adjusted continuously but are changed by a discrete amount when an adjustment takes place, which corresponds to our usual intuition about labor market institutions. The implied statistical model says that the true wage (or earnings) is constant for an unspecified period of time and then changes discretely at unspecified breakpoints. Thus, true wage changes in a noisy series can be estimated as one would estimate structural break dates in a standard time series. The Bai-Perron-Gottschalk method is to test for a structural break at all

[1]Surveys in some other countries have access to administrative data from payroll or tax records, which reduces measurement error significantly.

possible dates in a series. If one can reject the null hypothesis of no break for the most likely break date, then assume that there is a break at that point in time. Then examine the remaining sub-periods for evidence of structural breaks, and continue until one cannot reject the hypothesis of no break for all remaining dates. The adjusted series have wage (earnings) changes at all dates where we can reject the no-break hypothesis, and are constant otherwise. This is a systematic way of excluding many instances of transitory wage changes that look very much like measurement error. We present some examples in the paper where we apply this method to SIPP data for individuals in our sample.

We find the following main results. First, after correcting for measurement error, wages appear to be very sticky. In our baseline result with hourly workers, we find that the probability of a wage change is about 18 percent per quarter, thus implying an expected duration of wage contracts of 5.6 quarters. For salaried workers, the probability of an earnings change is only 5 percent per quarter. By comparison, several key papers estimating DSGE models using macro data estimate this probability to be about 30 percent per quarter, so we are finding significantly more nominal wage stickiness than macroeconomists typically assume. Second, the frequency of wage adjustment does not display any significant monthly or seasonal pattern. Third, we find little heterogeneity in the frequency of wage adjustment across industries. Wages in manufacturing appear to be stickier than wages in services. Similarly, we find little heterogeneity across occupations, although the frequency of wage adjustment is somewhat larger for service-related occupations. Fourth, we find that wage changes are significantly right-skewed, in keeping with preceding papers (for example, Gottschalk, 2005) that have found evidence of downward nominal wage rigidity in microdata. Fifth, the hazard of a nominal wage change first increases and then decreases, with a peak at 12 months. Thus, at a micro level, the pattern of wage changes appears somewhat more in keeping with the staggered contracting model of Taylor (1980) than with the constant-hazard model of Calvo (1983). However, our second result suggests that the timing of wage contracts is uniformly staggered throughout the year, which is the pattern that gives maximum persistence of nominal wages following a shock. Sixth, the probability of a wage change is positively correlated with the unemployment rate and with the rate of consumer price inflation. Finally, we show that higher wage stickiness makes it easier for macroeconomic models to match the stylized fact that monetary shocks cause persistent changes in real output and small but relatively persistent changes in prices.

This paper is connected to several strands of the literature. The first is the literature assessing wage rigidity using micro data. Much of this previous literature has concentrated on the different issue of downward nominal wage rigidity, rather than the frequency of wage adjustment *per se*. For example, Kahn (1997) reports evidence of a substantial downward nominal wage flexibility using U.S. data from the PSID. Gottschalk (2005) analyses data from the SIPP and, after correcting for measurement error, finds much greater downward wage rigidity. Lebow et al. (1999) use the Employment Cost Index and also find a substantial amount of downward wage rigidity. Dickens et al. (2007a) summarize the results coming from the International Wage Flexibility Project (IWFP).[2] The IWFP has collected data on wages and analyzed wage dynamics using data from a large number of countries. The main focus of this project is to perform analysis of wage dynamics that are comparable across different countries.[3] Perhaps unsurprisingly, one of the main findings of the project

[2] Other contributions are Dickens et al. (2007b) and Druant et al. (2008).

[3] The United States was included among the countries studied for assessing downward wage rigidity but not among those used to analyze wage stickiness. The reason, again, is that the IWFP data on U.S. wages comes from the PSID, which provides only annual data.

is that wage rigidity varies substantially across the different countries studied. This finding suggests that one should be careful in extrapolating our results to different countries and perhaps even to different time periods. Finally, in a very recent contribution, Heckel et al. (2008) analyze the frequency of wage adjustment for a large sample of French firms for the period 1998–2005.

Our paper is also related to the macro literature on nominal wage rigidity. Recent medium-scale macroeconomic models have used the sticky-wage assumption extensively. Most of these models, estimated through Bayesian techniques using aggregate data, suggest that nominal wages are quite sticky. However, as recently pointed out by Del Negro and Schorfheide (2008), this approach to estimation often delivers estimates that mirror the priors. In their conclusions, Del Negro and Schorfheide advocate more empirical analysis of microdata, along the lines of the work by Bils and Klenow (2004) and Nakamura and Steinsson (2008) on the frequency of price adjustment.[4] We view our paper as a first step towards providing similar micro estimates for wage dynamics.

A prominent strand of the literature on wage and employment dynamics over the business cycle has focused on search and matching models of the labor market.[5] Our paper is not directly related to this line of work. First, these papers are formulated in purely real terms, so the relevant concept is real wage rigidity, rather than the nominal rigidity we examine. Second, the search and matching framework indicates that the issue that matters for macroeconomic purposes is whether a pre-set wage paid to current employees is also applied to new hires.[6] We estimate the wage stickiness that matters in a monopolistically

[4] Although they warn that aggregation is a key issue when inferring macro behavior from micro evidence.
[5] For example, Shimer (2005) and Hall (2005).
[6] See Gertler and Trigari (2009).

competitive labor market setting, which is the frequency of wage changes for an individual regardless of employment history. Haefke et al. (2008) and Hall and Krueger (2008) examine micro evidence more related to the key predictions of the search literature.

Finally, our results shed some light on a small but interesting literature on the seasonal effects of monetary policy shocks. Recently, Olivei and Tenreyro (2008) have found that monetary policy shocks that occur in the first half of the year have larger real effects than those that occur later in the year. They explain this result by positing a model where wage changes are more likely to occur in the second half of the year. We find that while the frequency of wage changes is indeed slightly higher in the second half of the year, the magnitude of the difference is much smaller than assumed in the calibrated model of Olivei and Tenreyro, suggesting that a different model might be needed to explain their very interesting empirical finding.

The structure of the paper is as follows. The next section discusses the SIPP sample and the data definitions that we use. Section 3.3 summarizes the methodology we use to correct the wage series for unobserved measurement error. Section 3.4 contains the main results of the analysis obtained using the sample of hourly workers, while Section 3.5 presents the results for salaried workers. Section 3.6 explores the implications of the results obtained for the characteristics of a standard macroeconomic model. Section 3.7 contains the discussion of the hazard estimates and Section 3.8 concludes, and suggests some directions for future research.

3.2 Data

The data source for this paper is the Survey of Income and Program Participation (SIPP). The SIPP data have been collected by the Bureau of Labor Statistics (BLS) since 1983, with a major revision in 1996. The SIPP sample is a multi-stage, stratified, representative sample of the U.S. population. A large number of individuals are interviewed in order to collect detailed data regarding the source and amount of their income, a variety of demographic characteristics, and their eligibility for different federal programs. Each individual is followed for a period ranging from 24 to 48 months, with interviews taking place every four months.

The SIPP has at least two advantages compared to the other two large surveys used for this kind of analysis, namely the Outgoing Rotation Group (ORG) data from linked Current Population Surveys, and the Panel Study of Income Dynamics (PSID). First, unlike the PSID, the SIPP provides us with high-frequency information about wage changes. The near-quarterly frequency of the SIPP data makes it much more relevant for analyzing business cycles. Second, unlike the ORG, where an individual is interviewed for four consecutive months, not interviewed for the next eight months, and then interviewed for another four months before being dropped from the sample, the 1996 panel of the SIPP, which we use, follows each individual for up to 48 months, thus creating the proper panel data essential for our analysis.[7] Finally, the SIPP reports more-reliable information on wages and hours than the ORG or the PSID. In both these surveys, respondents are asked about their income only once a year, and must recall the amount and type of their income from various sources

[7]In fact, the CPS is even less suitable than this summary indicates, because the sampling unit is the household and not the individual. An individual leaving the housing unit is not followed; instead, new residents become survey members.

over the preceding calendar year – a daunting prospect for most people. By contrast, in the SIPP workers paid by the hour are asked specifically for their hourly wage rate in each interview.

We focus on the longest panel of this survey for which complete data are available: the 1996 panel (run from March 1996 to February 2000). For each person in the panel, we have time-series information about their wage rate (if they get paid by the hour) and monthly earnings as well as their industry and occupation. Table 3.1 reports basic descriptive statistics for our sample. The 1996 Panel follows 39,095 people, 49.4 percent of whom are women. We restrict our sample to workers between 15 and 64 years of age. The average person in our sample is around 38 years old.

Our first step aimed at minimizing measurement error is to focus on the smaller sample of those people who directly reported their hourly wage to the SIPP interviewer (because they are paid by the hour). Focusing only on hourly workers, however, has the drawback of reducing the size and representativeness of the sample for the full population. Thus, we also report separately the results for salaried workers. Of course, the latter do not report their wage rates. For salaried workers, we use monthly earnings as opposed to hourly earnings. The reason for this is that using hourly earnings would introduce much more noise into the wage history for salaried workers (since the earnings are divided by the worker's recollection of hours worked, which is also likely to suffer from measurement error).

Our smaller sample of hourly workers includes 17,148 people. Table 1 gives basic descriptive statistics of reported wages and earnings. The average wage rate in our sample for hourly workers is \$10.03. There is, however, a great deal of heterogeneity. The 5^{th} percentile of the distribution of wages is \$5 and the 95^{th} is \$20. The average monthly earnings for salaried workers is about \$3,000 dollars. The 5^{th} percentile of the distribution of earnings is \$440 and the 95^{th} is \$6,800.

Tables 3.2 and 3.3 report the breakdowns by industry and occupational category at the one-digit level. As Table 1 shows, services is the most highly represented industry (33.3 percent of total hourly workers), followed by trade (26 percent) and manufacturing (21 percent). Agriculture and mining, on the other hand, have very few observations. As for occupational categories, among hourly workers technical sales and support is the most highly represented in our sample (30 percent) followed by machine operators (24 percent) and services (19 percent). On the other hand, professionals and managers account for only 13 percent of the total in the hourly workers sample, while they represent almost 30 percent of the entire survey. Not surprisingly, our smaller sample under-represents occupational categories where workers are less likely to report receiving hourly wages.

3.3 Method

A key to our results is the need to limit the impact of measurement error in assessing the frequency of wage adjustment. Our first way of achieving this objective is to reduce the sample to the people who are hourly workers and reported their base wage rates to the SIPP interviewer. We prefer to concentrate on the hourly wage, since earnings may vary at the same wage if people change their hours worked. We also believe that people paid by the hour are likely to remember their hourly wage rates, but few people recall their monthly earnings down to the last dollar.

A second step is to apply to the reported data the procedure introduced by Gottschalk

(2005), which is intended to purge the wage series of unobserved measurement error.[8]

The procedure relies upon the Bai and Perron (1998 and 2003) method to test for structural breaks in the time-series context. The key identifying assumption is that wage changes take place in discrete steps. Assume that an individual works for T periods experiencing s wage changes at times $T_1...T_s$. The observed wage at time t, w_t, is equal to a constant α_t plus the unobserved measurement error ϵ_t:

$$
\begin{aligned}
w_t &= \alpha_1 + \epsilon_t \qquad t = 1...T_1 \\
&= \alpha_2 + \epsilon_t \qquad t = T_1...T_2 \\
&= ... \\
&= \alpha_{s+1} + \epsilon_t \qquad t = T_s...T.
\end{aligned}
\tag{3.1}
$$

The objective is to estimate the s break dates and the $s+1$ constant wages. The method proposed by Bai and Perron proceeds sequentially. First, using the whole sample of T observations, assume that there is one structural break, and pick the break date that minimizes the sum of squared residuals (SSR). Then test to see whether one can reject the null hypothesis of no break over the entire sample against the alternative that there is a break at the point that minimizes the SSR.[9] If one cannot reject the null, then the procedure is finished, and one concludes that there are no structural breaks in the sample (that is, the

[8] We also apply this procedure to earnings. We apply the same procedure to the reported wage series for hourly workers and to the reported earnings series for salaried workers. We refer only to wages in this section for expositional simplicity.

[9] Given the short length of the wage histories, the critical values for the structural break tests are obtained through Monte Carlo simulations. The appendix contains a more extensive explanation of the procedure adopted to compute the correct critical values.

wage is constant over the whole sample). If one can reject, then test for structural breaks in each of the subperiods identified by the break test. Again, pick the date that minimizes the SSR in each subperiod, and then test whether a significant break is detected at that point. Continue until no significant structural break is detected in any of the remaining subintervals of data.

One might object that this procedure is biased towards finding wages that are sticky, since the identifying assumption is that wages are set in nominal terms for a certain period of time! However, we do not constrain the procedure to assume any minimum number of periods between true wage changes. So, for instance, if an individual is followed for 48 months, corresponding to 12 interviews, the procedure can detect up to 10 wage changes.[10] One might then ask whether the procedure would be able to estimate a large number of breaks in a short time series. This is the important issue of the power of the test, which we discuss at length below.

Individual examples illustrate how this procedure works.[11] Figure 3.1 shows the reported and the adjusted wage series for "Linda," a 40-year-old secretary with high school degree. The reported series,[12] shown by the dashed line, is characterized by five wage increases and three wage decreases over the period considered. By contrast, the adjusted series, shown by the solid line, shows only two breaks, from \$12.54 to \$12.83 and from \$12.83 to \$13.56 (the last figure being the average of the subsequent reported wages).

Figure 3.2 reports the distribution of the measurement error as implied by our correc-

[10]Given that some people are observed for less than 48 months, we calculate a maximum quarterly frequency of true wage changes potentially obtainable. For people interviewed 12 times, for example, the probability is 83 percent (=10/12). Computing a weighted average of these probabilities across different numbers of interviews, we get a maximum detectable quarterly probability of a true wage change of 56 percent, which is much higher than we actually estimate

[11]We made up the names of the individuals in these examples.

[12](\$12.53, \$12.55, \$12.53, \$12.83,\$12.83,\$12.83, \$13.5 \$13.61 \$13.4 \$13.7 \$13.61)

tion procedure. As the figure shows, the distribution of the measurement error appears symmetric, with a big spike at zero.

Figure 3.3 shows the reported and the adjusted series for the hourly earnings of "Christina," a 40 year-old health care worker. Figure 3 highlights a potential problem with our methodology: the low power of the test for structural breaks (especially in a short time series) might lead us to underestimate the probability of a wage change. If people in our sample with true wage changes have sufficiently noisy reported wages (because of measurement error), then our test might fail to reject the null of no wage change.

Our application of trend-break tests to estimating the average frequency of wage changes raises several statistical issues. We present an in-depth discussion of these issues and our solutions in the appendix. Here we sketch an overview of this material.

The first problem is that standard F-tests are not the appropriate method to test for a structural break in wage series, for two conceptually different reasons. First, even in a setting where the error term is i.i.d. and Gaussian, a standard F-test can only be used to test a hypothesis about a structural change at a given date. By contrast, we perform tests for structural breaks at every possible date and then test for a structural break at the date where such a break appears most likely, and the distribution of this maximum F-statistic is different from a standard F distribution. This problem was addressed by Bai and Perron (1998). Second, the error term is not i.i.d., because the measurement error in wages is not a classical white noise error. In order to address this second problem, we need to know the structure of the measurement error. Fortunately, Gottschalk and Huynh (forthcoming) provide the required statistics. They compare reported SIPP wages with wages for the same people in an administrative data set, where wages are based on tax records and thus measured essentially without error. The difference between true and reported wages allows them to estimate the measurement error process for the SIPP. Armed with this information, we address the first and second problems together via Monte Carlo simulations (microsimulations). We simulate wage histories for individuals whose wages are constant in the simulations, but are assumed to be observed with measurement error of the type found by Gottschalk and Huynh. Based on these simulations, we calculate the critical value for the test of a structural break that will give us the required size of the test (the probability of Type I error).

Following this procedure ensures that we have a consistent estimator of the break dates for individual wage histories. However, just tabulating the frequency of these breaks does not give us a consistent estimator of the frequency of nominal wage changes in the population. Again, there are two reasons. The first is the size of the test. Even if all wages were constant, the fact that we pick the critical values to ensure a certain probability of Type I error means we would falsely conclude that α percent of wages change each period, where α is the size of the test. By itself, this force would lead us to conclude that wages are more flexible than is really the case. But of course there is also the issue of power — we do not necessarily detect every true break in the wage series. Low power will lead us to make the opposite mistake, and conclude that wages are more sticky than is really the case. We again address the issue of power and its implications for our estimated wage change frequency via simulation. Using simulations where true wages change in the sample, we can calculate the power of the test. Once the power of the test is known (of course, the size is also known, as it is a parameter we specify), we can adjust the raw frequency count of estimated breaks for both Type I and Type II error, and get a consistent estimate of the actual frequency of nominal wage changes. (See the appendix for details.) It is the frequency adjusted for Type I and Type II errors that we report in the next section.

3.4 Main Results

As noted in the introduction, we face the difficult task of mapping the large set of outcomes in micro data into simple macro models. To guide our exercise, we stick as closely as possible to estimating key parameters for the labor market institutions assumed in macro models with nominal wage stickiness, although these institutions surely characterize only a subset of the rich heterogeneity of employer-employee relationships present in our micro data. In macro models of this type, each worker is assumed to be a monopolistically competitive entrepreneur, supplying a unique variety of labor and setting his or her own wage. An example is the behavior of an independent contractor, such as a plumber or electrician, who charges according to a "rate sheet" specifying the wage charged per hour. Such a worker may work at a number of different residences over the course of a day, thus being paid by several different "employers" in quick succession and experiencing a number of very short "employment spells." Or the contractor might work on a single, large project for several weeks or even months, which would show up in the data as a long employment spell. But the rigidity of the contractor's nominal wage depends on the frequency with which she or he revises the rate sheet. In this framework, the right statistic to examine is the frequency of nominal wage changes (rate sheet revisions) over the entire sample for which we have data, disregarding any job transitions as irrelevant. For this reason, all the results presented in this section refer to the entire wage history of each individual, regardless of his or her employment history.

While our data and analysis are at conducted with interviews taking place every four months, we report the results at a quarterly frequency for ease of comparison with the previous literature.[13] Table 3.4 reports the frequency of wage adjustment for hourly workers. The quarterly frequency of wage adjustment for reported wages is very high. In the typical quarter for the 1996 panel, 48.1 percent of people report a different wage than they reported in the previous interview. The situation changes radically when considering the adjusted series for wages. In the 1996 panel, in the typical quarter only 17.8 percent of hourly workers experience a change in their wage.

Under the assumption that the correction for measurement error is appropriate, we therefore find evidence of greater wage stickiness than previously found in estimated DSGE models using aggregate data for the U.S. economy. CEE (2005), for example, estimated the quarterly Calvo probability of wage adjustment to be 36 percent in their benchmark model. In SW (2007) the benchmark estimate of the same parameter is 26.2 percent. Our results are more consistent with the findings of Gottschalk (2005), who used the same method but analyzed a previous wave of SIPP data. Gottschalk does not report exactly the parameter we estimate, but computing the analogous statistic from his adjusted wage series gives a figure closer to ours (11 percent). The difference can be partly explained by the fact that Gottschalk analyzes a different time period (1986–1993), and that arguably the U.S. labor market became more flexible over time. Finally, in a recent contribution, Heckel et al. (HLM, 2008) found the average quarterly frequency of wage adjustment to be 35 percent for a large sample of French firms. However, HML have access only to firm-occupation data;

[13] We transform the results into quarterly results by multiplying by $\frac{3}{4}$.

therefore, they cannot test the frequency of wage change at the individual level nor correct for measurement error in the reported wages.[14]

We should note that macro estimates of the nominal wage parameter are not always estimated to be around 0.30. In fact, CEE (2005) estimate variants of their baseline model, in two of which (no habit formation, and low investment adjustment costs) the estimated degree of wage stickiness is substantially higher, and close to our micro results. Perhaps future work on estimated DSGE models should simply follow CEE's policy and use micro evidence like ours as a way of disciplining macro estimates based on aggregate data.[15]

3.4.1 Seasonality

A second question we explore regards the seasonality of the pattern of wage adjustment. Olivei and Tenreyro (2008) find that monetary shocks have much larger effects on output if they occur in the first half of the year than if they occur in the last two quarters. They explain their findings by proposing a model where wage adjustment is seasonal, and is much more likely to take place in the second half of the year. Their calibrated model assumes that 24 percent of annual wage changes occur in the first quarter, 2 percent in the second quarter, 32 percent in the third quarter, and 42 percent in the fourth quarter. However, they explain that this calibration is based on a small sample of New England firms because "there is no systematic empirical evidence pointing to particular values for the [quarterly wage change frequencies]." We can supply this evidence using direct observation for a representative sample of the U.S. economy.

[14]Moreover, the mean wage in a firm-occupation cell can change even if there is no change in the wage for any individual worker, if the composition of the individuals in the cell varies over time

[15]CEE (2005, p. 40) write "Our position is that a reasonable contract length is one that matches the duration of contracts found in survey evidence. In this respect, we follow the empirical literature on wage and price frictions."

Figure 3.4 illustrates the frequency of wage adjustment by month. The frequency of wage adjustment for both the reported series and the adjusted series does not display the kind of sizable seasonal pattern assumed in Olivei and Tenreyro's calibration.

In order to investigate more formally the seasonality in the frequency of wage adjustment, we regress the probability of wage adjustment for both the reported and the adjusted wage series on a set of quarterly dummies, where the excluded category is the frequency of wage changes in the first quarter.

Table 3.5 reports the results. The F-test of joint significance of the quarterly coefficients always rejects the hypothesis that they are all zero. The results qualitatively support the assumption that drives Olivei and Tenreyro's model: Wage changes, do in fact, appear to be more likely in the second half of the year, not at the beginning. However, the magnitude of the difference is much smaller than Olivei and Tenreyro's calibration assumes. In our data, the proportion of quarterly wage changes relative to the total number of wage changes in a year is 23.6 percent in the first quarter, 24.1 percent in the second quarter, 26.6 percent in the third quarter and 25.5 percent in the last quarter. Whether these small differences can explain the differential seasonal effects of monetary policy shocks is an open question, but we suspect that they cannot.[16]

Interestingly, Heckel et al. (2008), using French firm-level data, find evidence that the frequency of wage adjustment is highly seasonal, with a spike in the third quarter. As the authors emphasize, this finding might be due to a very specific institutional feature of the French labor market, where by law the minimum wage is updated each year in July. However, there is no such feature in the U.S. labor market. Anecdotal evidence, in fact,

[16]See Dupor and Han (2009) for further discussion and a test of the hypothesis advanced by Olivei and Tenreyro (2008).

suggests that in the United States wage changes indeed take place in January in some firms, but in other firms they occur at the hiring date of the worker. In still other firms, wage changes are implemented at the beginning of the fiscal year.[17]

3.4.2 Heterogeneity

Our access to micro data allows us to explore whether wage stickiness differs across sectors or occupations.

Table 3.6 reports the results from regressing the probability of a wage change for hourly workers on a full set of industry dummies. The first two columns report the probability of a wage change obtained using, respectively, reported and adjusted wages. The third and the fourth columns report the differences relative to the manufacturing industry. While in general the hypothesis of total absence of heterogeneity is always rejected by the data, as shown by the p-value of the hypothesis that all the coefficients are zero, there is not much evidence of heterogeneity across industries in the frequency of wage changes. Services, trade and transport and communication display lower levels of wage skickiness than manufacturing, while construction, mining, and agriculture do not display significantly different coefficients.[18]

Table 3.7 repeats the exercise of Table 6, but now for different occupations. The coefficients in the third and fourth columns are relative to production workers. Again, we find only few significant results. One such result is that there is less wage stickiness in occupations related to services and in managerial occupation.

[17] In some peculiar cases, the wage changes take place on the birthday of the company!

[18] But this is also partly due to the imprecision of these estimates resulting from the very small number of observations for these industries.

The result of relatively little heterogeneity across industries and occupation might seem puzzling. We explored whether this is an artifact of aggregation by exploring the data at the 2-digits level of disaggregation (for both industries and occupations). While the point estimates display significantly more dispersion using the more refined classifications, the very small number of observation available for each category prevented us from finding differences that are statistically significant.[19]

3.4.3 Downward Nominal Wage Rigidity

In order to address a question typically asked by the labor literature on wage stickiness, we provide some evidence on the importance of downward nominal wage rigidities. Figure 3.5 reports the histogram of the non-zero adjusted wage changes. In order to avoid including outliers in the calculations, we plot only the inner 98 percentiles of the distributions (that is to say, we exclude the lowest and the highest percentiles). As the graphs show, wage reductions are much less frequent than wage increases. More precisely, they correspond to 11.5 percent of the non-zero wage changes. It is important to remember that we are analyzing one of the highest-growth periods of the last several decades, when nominal wage declines were probably less likely than in normal times.

Our results show that the period between 1996 and 1999 has been characterized by infrequent nominal wage cuts, which is normally taken as evidence of "downward nominal wage rigidity" in the literature. In our view, the term "rigidity" implies a friction, in this case a barrier to wages tracking the value marginal product of labor. Many authors have advanced the hypothesis that this barrier is stronger when the marginal product declines

[19]The results (not shown) are available upon request.

than when it rises. In our view, the evidence that we have produced (like the similar evidence that has been offered elsewhere in the literature) does not establish the existence of a rigidity in this sense. We think that a rigorous test of the hypothesis would require a model and further empirical work that would establish a baseline for the expected shape of the wage distribution. Absent such a baseline, we can establish that nominal wage cuts are infrequent, but this finding does not have a structural interpretation.

3.4.4 Cyclicality

A fourth question of interest is the correlation between the frequency of wage adjustment for hourly workers and some macroeconomic variables.

Figure 3.6 shows a scatter–plot of the cyclical components of the hp-filtered frequency of wage adjustment and the realized monthly inflation rate, computed as the percentage change of the U.S. CPI for all goods and all cities. As the figure shows, the two variables are positively correlated. The correlation between the two variables is 0.38 and is statistically significant.

Figure 3.7 shows a scatter-plot of the cyclical components of the hp-filtered frequency of wage adjustment against the monthly total U.S. unemployment rate. The relationship between the two variables is positive, with a correlation coefficient of 0.28, but it is not statistically significant .

In considering these results it is important to realize that the answers we provide in this section are based on an extremely short time-period (only four years of data), and hence must be considered with great caution. Subject to this caveat, the evidence suggests that pure time-dependent models of nominal wage rigidity, such as the models of Calvo and Taylor, cannot capture the whole story of infrequent nominal wage changes. Nominal wage adjustments appear to also have a state-dependent component. However, unlike the literature on nominal price rigidity, which has explored state-dependent models of price changes, we are not aware of models of nominal wage rigidity that are state- rather than time-dependent. Our results suggest that such models may be necessary. It would be useful to redo an exercise like ours using data from periods where the inflation and/or unemployment rates were high, to see whether the evidence of state-dependent wage changes is stronger in such periods.

3.5 Salaried Workers

While we chose to present our baseline results using only the sample of hourly workers for whom we have hourly wage data, here we also present estimates obtained using the sample of salaried workers. In order to do that, we focus on monthly earnings (as opposed to hourly earnings) and we apply the same measurement error correction procedure to earnings that we used for wages.

Table 3.8 reports the basic results for salaried workers. The average quarterly frequency of earnings change is about 65 percent. The results we get with the adjusted series are much smaller, 4.9 percent. In order to gain intuition about what might drive these extreme results, it is useful to discuss two examples.

Figure 3.8 shows the reported and adjusted series of monthly earnings for "Mark," a 42-year-old manager at a business service firm. This example clarifies how our procedure is able to capture any kind of change in earnings, provided that it is sufficiently large or persistent. Figure 9 reports instead the earnings history of “Peter,” a 39-year-old mechanic. The reported series is very noisy, with continual ups and downs in the reported earnings. Given the procedure explained in Section 3.3, the reader will not be surprised to see that in such a case the suggested adjusted earnings profile is totally flat. The predominance of reported earnings' histories similar to Peter's is the likely cause of the results shown in Table 3.8.

We repeat the same exercises we did for hourly workers on the sample of salaried workers. Figure 3.10 shows how the frequency of earnings changes varies by calendar month. It is hard to see any sizable seasonal effect. We investigate this hypothesis more formally, and Table 3.9 reports the results. While we always reject the hypothesis that wage changes are uniform over the year, the magnitudes of the differences are always very small.

The results for heterogeneity across industries and occupations are reported in Tables 3.10 and 3.11. While service and trade still exhibit higher wage flexibility than manufacturing, transport and communication industries, the group does not seem to have greater wage flexibility than manufacturing in the case of salaried workers. On the other hand, for salaried workers the only occupation category that displays a significant difference from the production workers category is that of service workers, who display more wage flexibility.

Finally, correlating the frequency of earnings adjustment for salaried worker with macro variables, we confirm a positive relationship with the unemployment and inflation rates, as in the case of hourly workers. The correlation with the monthly inflation rate is slightly lower (0.34). The correlation with the unemployment rate is 0.20 and is not statistically significant.

3.6 The Importance of Sticky Wages

We evaluate the significance of our findings for macroeconomics by using our parameter estimates in a benchmark medium-scale macro model. We use the DSGE model proposed by Smets and Wouters (2008; henceforth SW).[20] We take the model exactly as presented in their article, and we perform two simple exercises.

First, we estimate all the parameters of the model through Bayesian techniques after fixing the parameter for wage stickiness at 0.178 (our baseline result for hourly workers) and at 0.105.[21] Second, we compute the impulse response functions produced by the model following a monetary shock, using both the parameter estimates of the original SW paper and the estimates we obtained from our first exercise using data for hourly workers.

Table 3.12 reports the results obtained for some key parameters of the model. The first column reports the mode found in SW, the second column displays the posterior mode we find after fixing the wage stickiness parameter to 0.178, and the third column what we get with 0.105. As the table shows, the parameters related to price stickiness, price indexation, and wage indexation do not differ dramatically among the three cases. Also, the elasticity of intertemporal substitution appears to change very little. Interestingly, instead, the elasticity of labor supply increases a bit under our specification, going from 1.92 to 2.28 to 2.61. Finally, the capital share in the production function increases in our specification from the 0.19 obtained by SW to a more standard value of 0.3, which is also more consistent with long-run evidence from national income shares.

In terms of the dynamic responses to shocks, Figure 3.11 reports the impulse responses

[20] We use the code of the model that is available on the *American Economic Review* website.

[21] Roughly the weighted average of the findings obtained for hourly workers and salaried workers using the numbers of workers as weights.

to a monetary policy shock. The solid line is the impulse response from the SW model using the parameters reported in their paper, the long-dashed line reports the results from the model using our parameter estimates for hourly workers, and the dot-dashed line the results obtained for the totality of the workers. As expected, we find that with our estimates the model produces a larger and more persistent response of output and consumption to a monetary shock. This is not particularly surprising, since the micro data indicate that wage stickiness is higher than SW estimated based on aggregate data, while our estimates of the other structural parameters are substantially unchanged. The responses of hours do not differ dramatically, while the responses of the real wage and of price inflation appear to be damped and more persistent in our estimation.

Consistent with intuition, the findings show that higher wage stickiness makes it easier for macroeconomic models to match the stylized fact that monetary shocks cause persistent changes in real output and small but relatively persistent changes in prices.

3.7 Hazard Functions

Thus far, we have intentionally limited ourselves to computing a statistic that can be interpreted as the constant hazard of a wage (earnings) change, which is also the statistic estimated in macroeconometric models. However, our data allow us to test whether the hazard is truly constant, by estimating hazard functions. The estimated hazards let us compare the fit of Calvo-style models of wage rigidity—which imply a constant hazard of experiencing a wage (earnings) change—*vis a vis* other alternatives, such as contract renegotiations at fixed intervals as in the Taylor (1980) model, which would imply hazard functions that peak at certain durations.

To explore this issue we first use the reported and the adjusted wage (earnings) series to estimate a discrete-time hazard model, where an exit is defined as a change of the reported (or adjusted) wage (or earnings).[22] A new spell starts each time a new wage (or level of earnings) is observed, and we include in the sample all non-left-censored spells. We control for age, gender, and educational attainment, and we include a full set of duration dummies. We use the estimated coefficients on the duration dummies to find the hazard function for wage (earnings) changes.

Figure 3.12 shows the estimates of the hazard obtained using the reported wage series for the hourly workers and the reported earnings series for the salaried workers. The hazards are decreasing, with more than half of the respondents experiencing a wage change in the first four months. Declining hazards imply that the highest probability of having a wage change is immediately after the previous wage change. This pattern is intuitively unreasonable, suggesting that there is indeed significant measurement error in the reported wage. Figure 3.13 reports the estimates of the hazard obtained using the adjusted wage and earnings series. Here, by contrast, there is a clear peak at 12 months in both series.

We conclude that Taylor-type fixed-length contracts have stronger empirical support than Calvo-type constant-hazard models. However, the fact that the wage change frequency is almost flat over the calendar year (Figure 4 and Table 5) suggests that the starting time of the wage contracts is uniformly staggered throughout the year. This pattern is, of course, the one that gives the largest contract multiplier and creates maximum persistence of the real effects of nominal shocks. Although it gives the greatest persistence, uniform staggering

[22]See Box-Steffensmeier and Jones (2004) p. 73.

is typically found to be an unstable Nash equilibrium, so it is interesting that we are finding indirect evidence of staggered, rather than synchronized, wage contracts.[23]

3.8 Conclusion

Since we already outlined the main results in the introduction, we conclude by suggesting directions that future research might take.

First, it is important to understand why the stickiness estimated from micro data is greater than that estimated from aggregate data using Bayesian techniques. Idiosyncratic measurement error, such a large concern in the analysis of micro data, is unlikely to be the explanation. Such errors would average out and contribute little to the variance of any aggregate wage series. One possibility is that the difference is due purely to aggregation issues: for example, if high-wage workers' wages also adjust more frequently, then the aggregate wage will appear to more flexible than the average worker's wage. We plan to investigate this possibility using our data, but since high-wage workers are likely to be salaried workers, whose adjusted earnings we find to be stickier than the wages of hourly workers, this explanation appears unlikely. The reasons for this micro-macro gap should shed light on the perplexing issues of aggregation that must concern all macroeconomists interested in "structural" models.

Second, the lack of sizable seasonality in wage changes leaves an open question: what can explain the estimated differential effects of monetary shocks occurring in different quarters?

[23]Our findings are consistent with the empirical studies of Taylor (1983) and Cecchetti (1984), who found staggered wage setting in union contracts. However, in the U.S. labor market, very few workers are covered by formal union contracts, so it is useful to extend their results to a representative sample of the U.S. labor force. Some notable papers show that in richer models staggering might be a stable Nash equilibrium after all. See, for example, Fethke and Policano (1984), Ball and Cecchetti (1988), and Bhaskar (2002).

Nakamura and Steinsson's (2008) finding that price adjustment is seasonal suggests one answer.[24]

Third, the findings on the shape of the hazard functions suggest that we should explore the properties of models based on fixed-length wage contracts, as in Taylor (1980), in addition to the very tractable stochastic-length contracting models in the style of Calvo (1983).[25]

Fourth, our desire to estimate the key parameter of one particular macro-labor model led us to focus on wage histories and disregard employment histories. However, the implication that employment history is irrelevant is not shared by all macro models of the labor market. For example, in the literature on search and matching in business cycle models, the wage stickiness that matters for macro is the degree of (real) wage rigidity for new hires. We plan to explore further these issues in future research.

Finally, from an epistemological point of view, we hope that this work will increase the awareness that greater communication between economists working in different fields (in this case, macro and labor economics) can produce valuable insights at relatively low cost.

[24]Of course, Dupor and Han (2009) question the robustness of the finding itself.
[25]See, for instance, Knell (2010).

3.9 Technical Appendix

This appendix describes the methods we use to (1) obtain critical values to test for changes in wages when the assumptions for the standard F test are not met, (2) estimate the power of these tests, and (3) obtain unbiased estimates of the probability of a wage change by correcting for the impact of Type I and Type II errors.

3.9.1 Critical values

The standard F test cannot be used to test for wage changes since the necessary assumptions for the F tests are violated in two conceptually different ways. First, measurement error in earnings is not classical.[26] The critical value must, therefore, be adjusted to take account of this violation of the assumptions. Second, the test for structural breaks used in this paper is a test of the maximum of a set of F statistics rather than the test of a single F statistic. Bai and Perron (1998) show that the appropriate test for structural breaks must take into account that the test is based on the maximum of l test statistics, where l is the length of the period being analyzed. The standard critical values are no longer applicable since the critical value for the maximum of l test statistics is higher the critical value for a single F statistic.

We address these problems by using Monte Carlo simulations that simulate data with the same non-classical measurement structure found in the SIPP by Gottschalk and Huynh (forthcoming). Their estimate of the structure of measurement error in the SIPP was obtained from SIPP earnings records matched to uncapped W-2 earnings records in the Detailed Earnings Records (DER) file. Measurement error is defined as the difference between DER earnings and reported earnings in the SIPP. These matched records are used to estimate the autocorrelation of measurement error (0.54) and the signal-to-noise ratio (2.64).[27]

In order to obtain critical values to test the null hypothesis of no change in wages, we generate 2,000 wage profiles of length l with no change in wages.[28] We then apply the method described in the paper to test for structural breaks in each of these constant wage series with measurement error. The critical value for a test with a significance level of α is obtained by calculating the F value for each wage series, ranking these wage series on the basis of their F values, and finding the critical F value where we falsely reject the null of no changes in wages α percent of the time. This is repeated for simulated earnings series of length $l = 3, 4...L$

[26] See the large literature reviewed in Bound and Mathiowetz (2001)

[27] Gottschalk and Huynh (forthcoming) also report a negative correlation between measurement error and DER wages of -0.339. Whether this mean reversion should be included in the analysis of individual wage profiles depends crucially on whether the negative correlation is between group (the expected value of measurement error is lower for respondents with above average earnings) or within group (the expected value of measurement error declines when an individual's wages rise). We assume the correlation is between group so mean reversion affects the mean but not the variance of reported wages and has no impact on our estimate of the probability of wage change.

[28] Including multiple changes in wages over the l periods would not affect the estimates since the algorithm in the first iteration is based on the maximum F statistic over the full l periods, no matter how many wage changes are found in further iterations.

3.9.2 Power

We obtain the power of the test using a similar simulation procedure. In order to obtain the power of the test we similarly generate 2,000 wage series of length l. But in this case each wage series has a permanent wage change of Δw after $\tilde{l} < l$ periods, where $\tilde{l}$ is randomly assigned. This is the true wage series. The observed wage series has measurement error around this non-constant wage series. The variance of the measurement error is set to be consistent with the signal-to-noise ratio found by Gottschalk and Huynh (forthcoming). These wage series are used to calculate the proportion of times our tests falsely fail to reject the null of no wage change, using the previously discussed critical values. This yields the power of the test for a wage history of length l using a significance level of α.

3.9.3 Adjustments to obtain consistent estimates

Estimating π, the probability that an individual will experience a wage change between waves, requires an adjustment to $\hat{\pi}$ for Type I and II errors. We start by defining an indicator variable, $I_{ip}(W_i, \alpha, \gamma|X)$, that takes the value 1 when the Bai and Perron (2003) algorithm finds a statistically significant change in person i's wages in period p, given the person's wage history, W_i, the significance level α used to test for breaks, and the power of this test, γ, given the number of observations over which the maximum F is calculated.[29]

The object of interest is the probability of a wage change across all persons and periods. Our sample analogue estimator is

$$\hat{\pi} = \frac{\sum\limits_{i=1}^{N} \sum\limits_{p=1}^{p_i^T} I_{ip}}{\sum\limits_{i} p_i^T}, \tag{3.2}$$

where p_i^T is the total number of periods in i's wage history. The numerator is the total number of statistically significant wage changes across the N histories, while the denominator is the total number of periods across all histories.

The Bai-Perron algorithm provides a consistent estimate of I_{ip}, the indicator of whether a wage change has occurred in period p of person $i's$ wage history. However, this does not ensure that $\hat{\pi}$ is a consistent estimator of π, since the tests for wage changes are subject to both Type I and Type II errors.

Consider conducting $P = \sum\limits_i p_i^T$ tests for structural breaks. In expectation, $P(1 - \pi)$ of these tests will be in periods where $\Delta w = 0$. However, as a result of Type I error, $\alpha P(1 - \pi)$ of the tested segments with no wage change will be falsely classified as having a statistically significant wage change. This error will lead us to over-estimate π. On the other hand, Type II error (failing to reject the null of no wage change when it is false), leads to an underestimate of π. The expected value of the number of wage changes that

[29]The X vector includes observables that affect the critical value, such as the length of the period over which the test is being carried out. X may also include variables of substantive interest, such as occupation or cyclical variables. In practice, we condition on these probabilities by including a set of dummies in descriptive logit estimates. These include a set of dummies for the length of the period used to calculate the maximum F.

are falsely classified as having constant wages due to sampling error is $(1-\gamma)P\pi$, where γ is the power of the test.[30]

The net impact of Type I and Type II errors is[31]

$$p\lim(\hat{\pi}) = \frac{\alpha P(1-\pi)+\gamma P\pi}{P} \quad (3.3)$$
$$= \alpha(1-\pi)+\gamma\pi \quad (3.4)$$
$$= \alpha+(\gamma-\alpha)\,\pi, \quad (3.5)$$

which implies

$$p\lim\left[\frac{\hat{\pi}-\alpha}{(\gamma-\alpha)}\right]=\pi. \quad (3.6)$$

Let $\tilde{\pi}=\frac{\hat{\pi}-\alpha}{(\gamma-\alpha)}$, so $\tilde{\pi}$ is a consistent estimator of the probability that a tested wage change is non-zero. It can be computed by using our initial estimate of the probability of a wage change, $\hat{\pi}$, the significance level used in these tests, α, and the power of these tests, γ, that we obtain from the microsimulations.

[30] Power is defined as one minus the probability of a Type II error. Power depends on the significance level of the test for a wage change and the number of periods used in the estimation.

[31] This adjustment for Type I and Type II errors would seem to be applicable to a wider set of estimators in which $\hat{\theta}$ is function of a set of estimators $\hat{\gamma}_j(x)$ from a lower level of aggregation, j, each of which is subject to Type I and Type II errors. Estimators using imputed values are one such example.

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Table 3.1: **Descriptive Statistics**

Total People (beginning)	39,095
Females	19,321
Hourly Workers (beginning)	17,148
Females	8,931
Mean Age	38
Mean Wage (Hourly Workers)	$10.03
Mean Earnings (Salaried Workers)	$2942

Table 3.2: **Industry Composition of the Sample**

Sample	**Total**	**Total**	**Hourly**	**Hourly**
Agriculture	778	1.99 percent	386	2.25 percent
Mining	174	0.45 percent	73	0.43 percent
Construction	1,993	5.10 percent	1,128	6.58 percent
Manufacturing	6,785	17.36 percent	3,684	21.48 percent
Transport and Communication	2,736	7.00 percent	1,093	6.37 percent
Trade	8,168	20.89 percent	4,459	26.00 percent
Services	15,881	40.62 percent	5,721	33.36 percent
Government and Public Administration	2,377	6.08 percent	597	3.48 percent
Army and Unemployed	203	0.52 percent	7	0.04 percent
Total	**39,095**	**100 percent**	**17,148**	**100 percent**

Table 3.3: **Occupational Composition of the Sample**

Sample	**Total**	**Total**	**Hourly**	**Hourly**
Professional	5,660	14.48 percent	1,033	6.02 percent
Managerial	4,932	12.62 percent	638	3.72 percent
Technical Sales and Support	11,761	30.08 percent	5,109	29.79 percent
Craftsmen and production	4,048	10.35 percent	2,337	13.63 percent
Operatives	6,185	15.82 percent	4,239	24.72 percent
Service	5,504	14.08 percent	3,360	19.59 percent
Farming	807	2.06 percent	426	2.48 percent
Miscellaneous and Unemployed	198	0.51 percent	6	0.03 percent
Total	39,095	100 percent	17,148	100 percent

Table 3.4: **Quarterly Frequency of Wage Adjustment, Hourly Workers**

Period	**Reported**	**Adjusted**	**CEE 05**	**SW 07**	**HLM 08**	**Gottschalk 05**
96-99	0.481	0.178				
65-95			0.36			
66-04				0.26		
98-05					0.35	
86-93						0.11

Table 3.5: **Seasonality of the Frequency of Wage Adjustment, Hourly Workers, Results Relative to the First Quarter**

Type of Wages	**Reported**	**Adjusted**
Q_2	-0.002	0.003
	(0.003)	(0.003)
Q_3	0.020***	0.021***
	(0.003)	(0.003)
Q_4	0.012***	0.013***
	(0.003)	(0.003)
F-test	0.000	0.000
N	136043	136043

Table 3.6: **Heterogeneity in the Quarterly Frequency of Wage Adjustment, by Industry, Hourly Workers (excluded category: Manufacturing)**

Wages	**Reported Levels**	**Adjusted Levels**	**Reported Relative**	**Adjusted Relative**
Agriculture	0.423***	0.178***	-0.081***	0.010
	(0.008)	(0.009)	(0.008)	(0.009)
Mining	0.482***	0.148***	-0.022	-0.020
	(0.014)	(0.015)	(0.014)	(0.015)
Construction	0.456***	0.174***	-0.048***	0.007
	(0.004)	(0.004)	(0.004)	(0.005)
Manufacturing	0.504***	0.168***		
	(0.002)	(0.002)		
Transport and Communication	0.501***	0.186***	-0.003	0.018***
	(0.004)	(0.004)	(0.004)	(0.005)
Trade	0.475***	0.186***	-0.029***	0.018***
	(0.002)	(0.002)	(0.003)	(0.003)
Services	0.473***	0.180***	-0.031***	0.012***
	(0.002)	(0.002)	(0.003)	(0.003)
Govt and Pub. Admin.	0.511***	0.168***	0.007	-0.000
	(0.005)	(0.006)	(0.006)	(0.006)
F-test	0.000	0.000	0.000	0.000
N	136043	136043	136043	136043

SE in parenthesis. ***, **, * : Statistically significant at 1 percent, 5 percent and 10 percent

Table 3.7: **Heterogeneity in the Quarterly Frequency of Wage Adjustment, by Occupation, Hourly Workers (excluded category: Production Workers)**

Wages	**Reported Levels**	**Adjusted Levels**	**Reported Relative**	**Adjusted Relative**
Professional	0.476***	0.171***	-0.010**	-0.001
	(0.004)	(0.004)	(0.005)	(0.005)
Managerial	0.455***	0.186***	-0.031***	0.014**
	(0.005)	(0.006)	(0.006)	(0.006)
Technical Sales and Support	0.478***	0.184***	-0.008**	0.012***
	(0.002)	(0.002)	(0.003)	(0.004)
Craftsmen and production	0.486***	0.172***		
	(0.003)	(0.003)		
Operatives	0.494***	0.174***	0.008**	0.003
	(0.002)	(0.002)	(0.003)	(0.004)
Service	0.478***	0.180***	-0.008**	0.008**
	(0.002)	(0.003)	(0.003)	(0.004)
Farming	0.442***	0.182***	-0.044***	0.011
	(0.007)	(0.008)	(0.007)	(0.008)
F-test	0.000	0.000	0.000	0.000
N	136043	136043	136043	136043

SE in parenthesis. ***, **, * : Statistically significant at 1 percent, 5 percent and 10 percent

Table 3.8: **Quarterly Frequency of Earnings Adjustment, Salaried Workers**

Type	**Reported**	**Adjusted**
1996-1999	0.653	0.049

Table 3.9: **Seasonality of the Frequency of Earnings Adjustment, Salaried Workers, Results Relative to the First Quarter**

Type of Earnings	Reported	Adjusted
Q_2	-0.007***	0.001
	(0.002)	(0.001)
Q_3	0.012***	0.007***
	(0.002)	(0.001)
Q_4	0.008***	0.006***
	(0.002)	(0.001)
F-test	0.000	0.000
N	187694	185980

Table 3.10: **Heterogeneity in the Quarterly Frequency of Earnings Adjustment, by Industry, Salaried Workers (excluded category: Manufacturing)**

Earnings	Reported Levels	Adjusted Levels	Reported Relative	Adjusted Relative
Agriculture	0.622***	0.050***	-0.036***	0.003
	(0.005)	(0.003)	(0.005)	(0.003)
Mining	0.621***	0.047***	-0.037***	-0.000
	(0.008)	(0.004)	(0.008)	(0.004)
Construction	0.669***	0.053***	0.011***	0.006***
	(0.003)	(0.002)	(0.003)	(0.002)
Manufacturing	0.658***	0.047***		
	(0.002)	(0.001)		
Transport and Communication	0.669***	0.048***	0.011***	0.000
	(0.002)	(0.001)	(0.003)	(0.001)
Trade	0.665***	0.050***	0.007***	0.003***
	(0.001)	(0.001)	(0.002)	(0.001)
Services	0.645***	0.051***	-0.012***	0.004***
	(0.001)	(0.000)	(0.002)	(0.001)
Govt and Pub. Admin.	0.658***	0.047***	-0.000	-0.000
	(0.002)	(0.001)	(0.003)	(0.001)
F-test	0.000	0.000	0.000	0.000
N	187694	185980	187694	185980

SE in parenthesis. ***, **, * : Statistically significant at 1 percent, 5 percent and 10 percent

Table 3.11: **Heterogeneity in the Quarterly Frequency of Earnings Adjustment, by Occupation, Salaried Workers (excluded category: Production Workers)**

Salaried	**Reported Levels**	**Adjusted Levels**	**Reported Relative**	**Adjusted Relative**
Professional	0.637***	0.050***	-0.032***	0.002
	(0.001)	(0.001)	(0.002)	(0.001)
Managerial	0.639***	0.048***	-0.030***	-0.001
	(0.001)	(0.001)	(0.002)	(0.001)
Technical Sales and Support	0.657***	0.050***	-0.012***	0.002
	(0.001)	(0.001)	(0.002)	(0.001)
Craftsmen and production	0.669***	0.048***		
	(0.002)	(0.001)		
Operatives	0.684***	0.048***	0.015***	-0.000
	(0.002)	(0.001)	(0.003)	(0.001)
Service	0.673***	0.052***	0.004	0.004***
	(0.002)	(0.001)	(0.003)	(0.001)
Farming	0.633***	0.050***	-0.036***	0.002
	(0.005)	(0.003)	(0.005)	(0.003)
F-test	0.000	0.000	0.000	0.003
N	187694	185980	187694	185980

SE in parenthesis. ***, **, * : Statistically significant at 1 percent, 5 percent and 10 percent

Table 3.12: **SW (2008) DSGE Model: Estimated Structural Parameter Values Fixing Wage Stickiness (Selected Parameters, Posterior Mode)**

SW Parameter	**Meaning**	**SW Original**	**BBG Hourly**	**BBG Total**
ξ_w	Wage Stickiness	**0.262**	**0.178**	**0.105**
ξ_p	Price Stickiness	0.66	0.66	0.69
ι_w	Wage Indexation	0.59	0.57	0.55
ι_p	Price Indexation	0.23	0.21	0.20
σ_c	EIS	1.40	1.37	1.40
σ_l	Elast. of Labor Supply	1.92	2.28	2.61
α	Capital Share	0.19	0.30	0.30

Figure 3.1: **Adjusted Wage Series, An Hourly Worker**

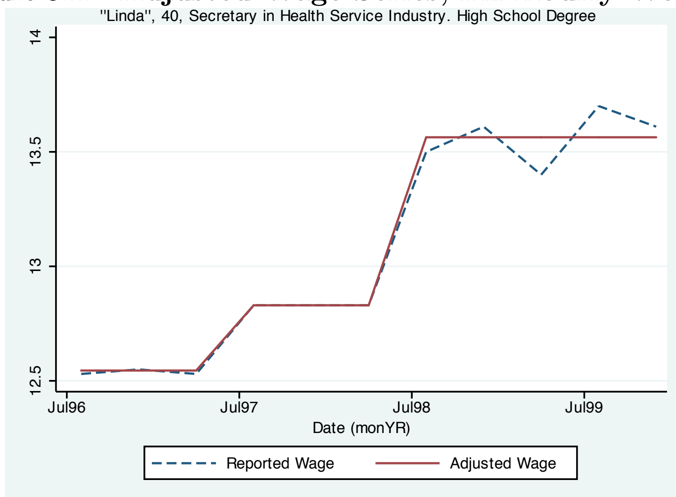

Figure 3.2: **Implied Measurement Error, Hourly Workers, 1996 Panel**

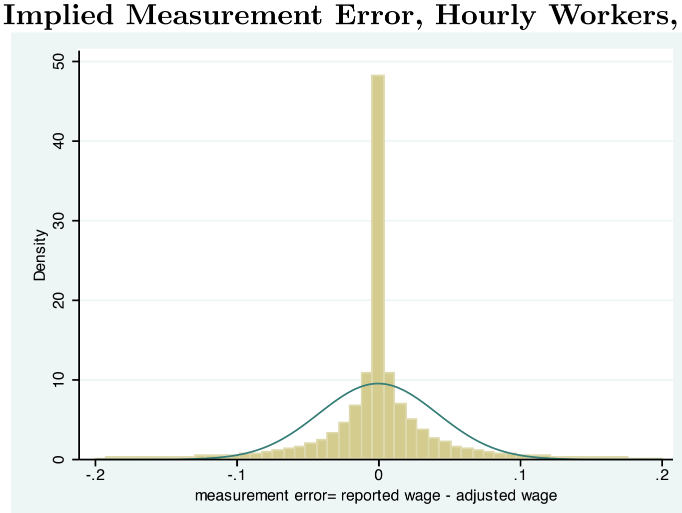

Figure 3.3: **Adjusted Wage Series, An Hourly Worker**

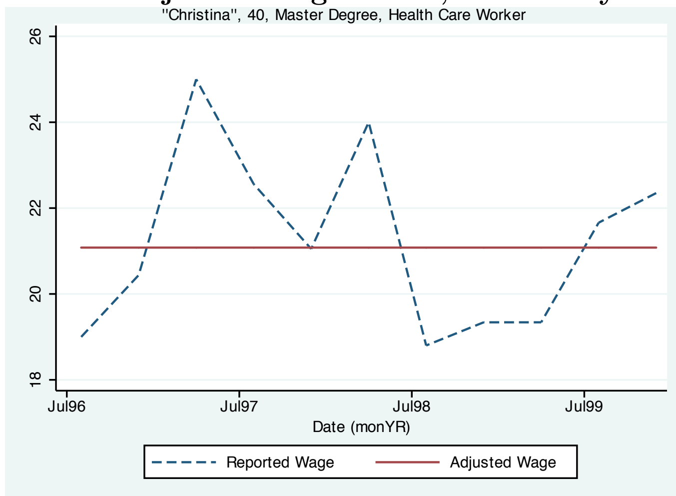

Figure 3.4: **Seasonality in the Frequency of Wage Adjustment, Hourly Workers**

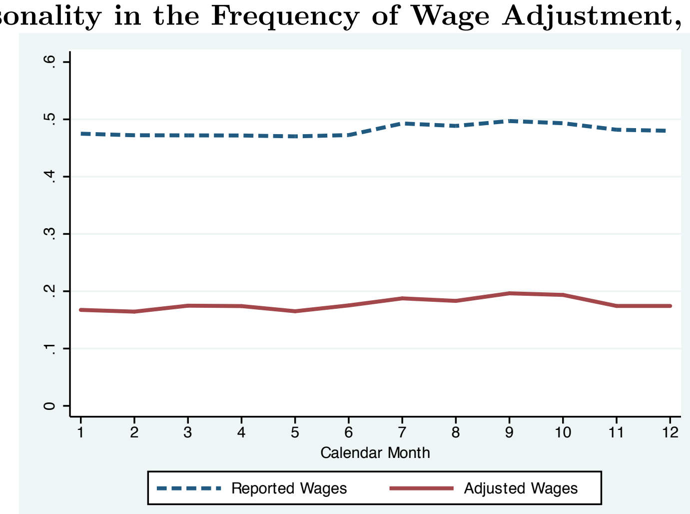

Figure 3.5: **Distribution of Non-Zero Wage Changes, Hourly Workers, 1996 Panel**

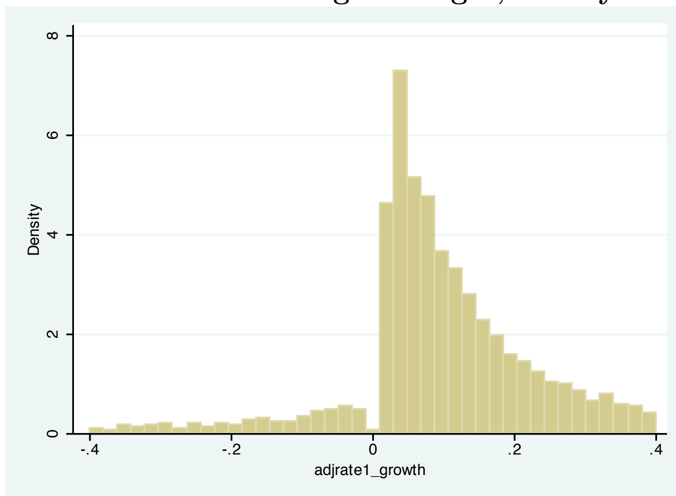

Figure 3.6: **Frequency of Wage Change and Monthly Inflation Rate, Hourly Workers, 1996-1999**

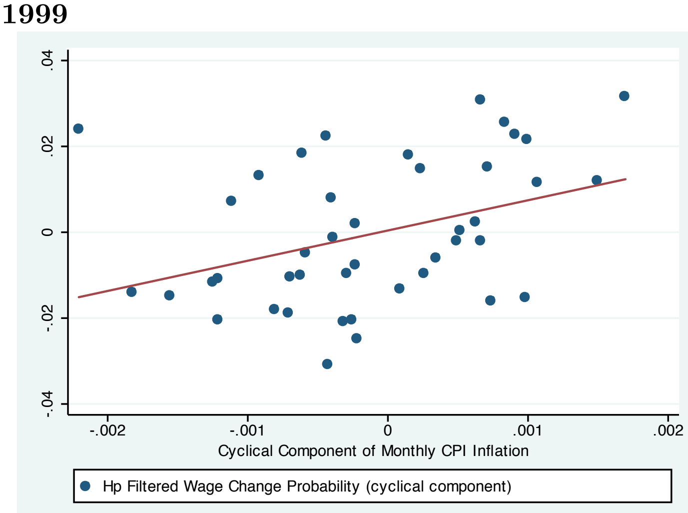

Figure 3.7: **Frequency of Wage Change and Unemployment Rate, Monthly, Hourly Workers, 1996-2003**

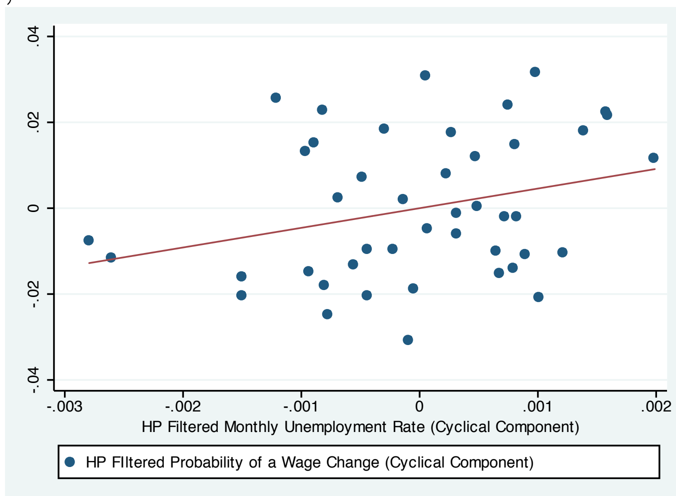

Figure 3.8: **Adjusted Wage Series, a Salaried Worker**

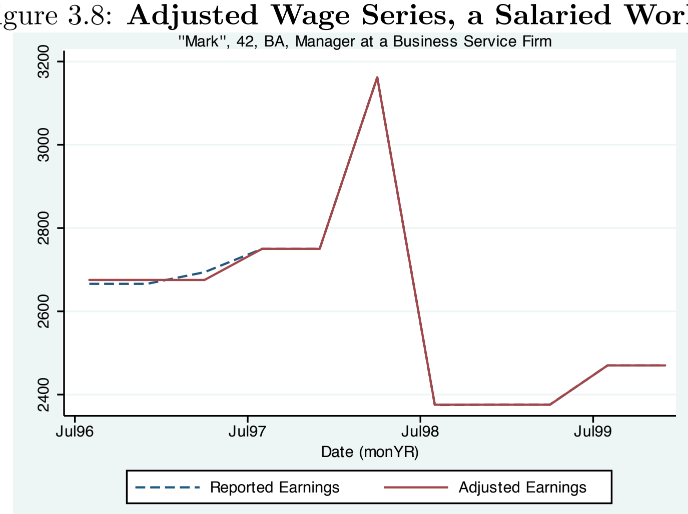

Figure 3.9: **Adjusted Wage Series, a Salaried Worker**

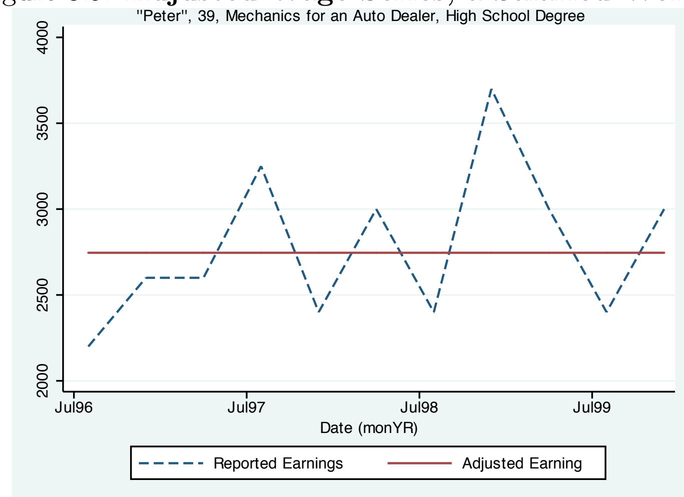

Figure 3.10: **Seasonality in the Frequency of Wage Adjustment, Salaried Workers**

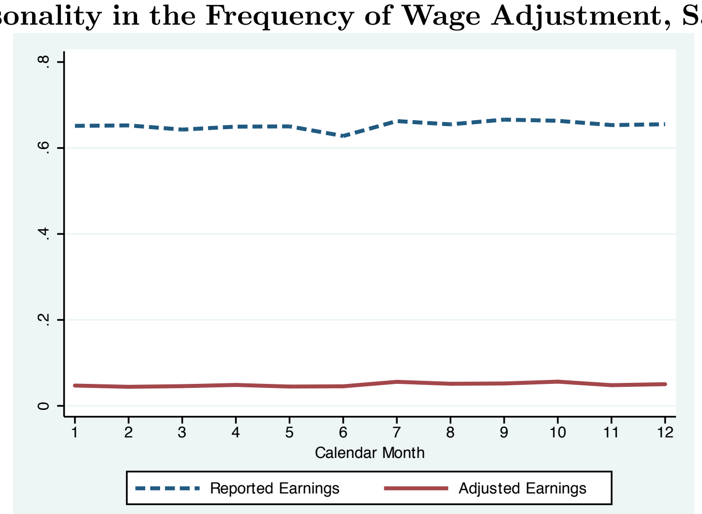

Figure 3.11: **Dynamic Response to a Monetary Shock, SW (2008) Model, Different Levels of Wage Stickiness**

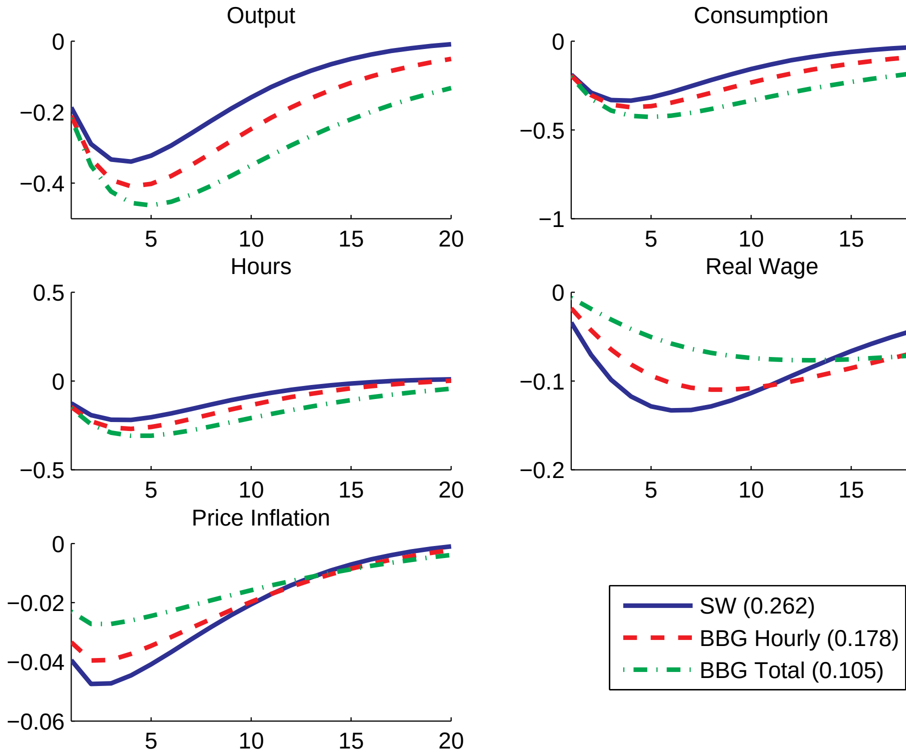

Figure 3.12: **Hazard of a Wage Change, Reported Wages, 1996 Panel**

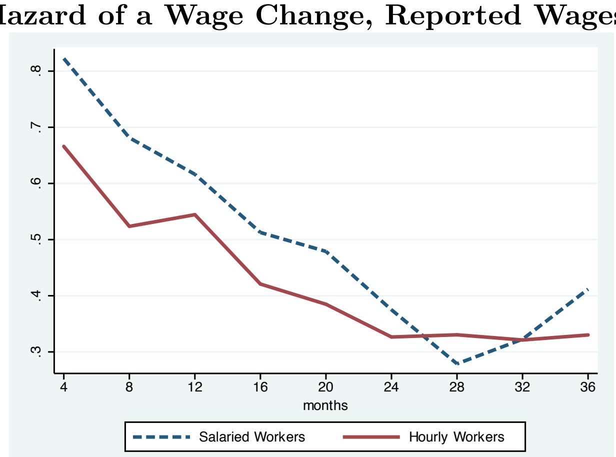

Figure 3.13: **Hazard of a Wage Change, Adjusted Wages, 1996 Panel**