# Essays on the Transmission of Economic Shocks


Claire H. Hollweg


A thesis submitted for the degree of
Doctor of Philosophy in
The School of Economics of
The University of Adelaide

May 2014

# Table of Contents

# Abstract

This thesis explores the transmission of economic shocks. Although the thesis is structured as four stand-alone chapters, the common theme throughout is identifying the impact of economic shocks: either idiosyncratic shocks at the household-level, macroeconomic shocks emanating from foreign countries and transmitted through global markets, or countries' own macroeconomic policy changes (for example, structural reforms or trade reforms). Each chapter applies a different empirical methodology, including structural estimation, reduced form instrumental variables estimation, and growth accounting. Finally, each chapter utilizes a different dataset and country sample selection. While one chapter uses a micro dataset from household-level surveys, others use cross-country datasets at the aggregate country level. Both developed and developing countries are considered in the analyses.

The thesis begins by exploring the relationship between idiosyncratic income changes and consumption changes of Australian households over the period 2001-2009. A major contribution to the literature is the use of the Household Income and Labor Dynamics of Australia dataset that includes panels on both consumption and income data. For the entire sample of Australian households, nearly full consumption smoothing exists against transitory shocks. Although less consumption smoothing exists against permanent shocks, Australian households still achieve a high degree of consumption smoothing against highly persistent shocks, particularly when compared to households in the United States. Durable purchases, female labor supply, and taxes and transfers are all found to act as consumption-smoothing mechanisms.

The thesis then explores the impact of structural reforms on a comprehensive list of macro-level labor-market outcomes, including the unemployment rate, employment levels, average wage index, and labor force participation rates. After documenting the average trends across countries in the labor-market outcomes up to ten years on either side of each country's reform

year, fixed-effects ordinary least squares as well as instrumental variables regressions are performed to account for likely endogeneity of structural reforms to labor-market outcomes. Overall the results suggest that structural reforms lead to positive outcomes for labor, particularly for informal workers. Redistributive effects in favor of workers, along the lines of the Stolper-Samuelson effect, may be at work.

The thesis then explores the impact of trade liberalization on macroeconomic estimates of productivity using Brazil as a case study. Trade and economic reforms can affect the price of capital goods relative to other tradable and especially non-tradable goods. If the price of capital investments rises more than the price of all goods and services in the economy, mis-measurement of the price of capital caused by the divergence in these relative prices would result in an overestimated capital stock and underestimated TFP. This chapter overcomes this bias by constructing a capital price index using international trade data on capital goods' unit values then adjusts the index to reflect domestic Brazilian prices. A significant recovery between 1992 and 2006 is observed, highlighting the important role of the price deflator in growth accounting.

The final chapter of this thesis proposes a methodology to measure the vulnerability of a country through exports to fluctuations in the economic activity of foreign markets. Export vulnerability depends first on the overall level of export exposure, measured as the share of exports to a foreign market in gross domestic product, and second on the sensitivity of exports to fluctuations in foreign gross domestic product. This sensitivity is captured by estimating origin-destination specific elasticities of exports with respect to changes in foreign gross domestic product using a gravity model of trade. Although the results suggest differences in elasticity estimates across regions as well as product categories, the principal source of international heterogeneity in export vulnerability results from differences in export exposure to global markets.

# Declaration

I certify that this work contains no material that has been accepted for the award of any other degree or diploma in my name, in any university or other tertiary institution and, to the best of my knowledge and belief, contains no material previously published or written by another person, except where due reference has been made in the text. In addition, I certify that no part of this work will, in the future, be used in a submission in my name, for any other degree or diploma in any university or other tertiary institution without the prior approval of the University of Adelaide and where applicable, any partner institution responsible for the joint-award of this degree.

I give consent to this copy of my thesis, when deposed in the University Library, being made available for loan and photocopying, subject to the provisions of the Copyright Act 1968.

The author acknowledges that copyright of published works contained within the thesis (as listed below) resides in the copyright holder(s) of those workers:

- Hollweg, C. H., D. Lederman, and J.-D. Reyes (2012). "Monitoring Export Vulnerability to Changes in Growth Rates of Major Global Markets." Police Research Working Paper Series No. 6266, The World Bank, Washington, D.C.

I also give permission for the digital version of this thesis to be made available on the web, via the University's digital research repository, the Library Search and also through web search engines, unless permission has been granted by the University to restrict access for a period of time.

Claire H. Hollweg, 1 May 2014

## Acknowledgements

This thesis would not have been possible without the support given by numerous people and organizations.

My sincere acknowledgements and thanks go to Associate Professor Markus Brückner, Dr. Daniel Lederman, and Professor Kym Anderson for helpful advice, comments, and supervision over the past years. I am also grateful to Professor Christopher Findlay and Dr. Jacob Wong for their continual encouragement and guidance throughout my postgraduate studies. I also thank two anonymous examiners for their insightful comments and suggestions.

On the financial side, I am in gratitude for support from the International Trade Department of the World Bank, the School of Economics of the University of Adelaide, as well as the University of Adelaide Graduate Centre for funding provided by the Adelaide Graduate Fee Scholarship.

On a personal note I would like to thank my husband Kevin Jones, my sister Terill Hollweg, and my parents Leslie and Joseph Hollweg for their continued love and support.

## Structure of Thesis

This thesis contains four chapters, which are stand-alone pieces with self-contained references, tables and figures.

The first chapter, titled "Measuring Consumption Smoothing in Australia: An empirical test of the permanent income hypothesis", uses microeconomic data to estimate a structural macroeconomic model of the degree of consumption smoothing to income shocks of Australian households. This chapter is a replication of a study undertaken in the United States.

The second and third chapters study the impact at the country level of government-imposed permanent shocks. The second chapter, titled "Structural Reforms and Labor Market Outcomes: International panel data evidence", is a cross-country empirical analysis of the impact of trade-related and other structural reform shocks on labor-market outcomes. This chapter has been used as a background paper into the report titled "Sticky Feet: How Labor Market Frictions Shape the Impact of International Integration on Labor Market Outcomes" by C. H. Hollweg, D. Lederman, D. Rojas, and E. Ruppert Bulmer. A version of this chapter has also been submitted for publication to *The World Economy*.

The third chapter, titled "Measuring Capital Matters", is a case study for Brazil and considers the impact of trade liberalization on the relative price of capital and resulting productivity measures. This chapter has been used as a background paper into the report titled "Brazil's Productivity Challenge after Economic Reforms" by R. Clark, L. De Zoratto, M. Dutz, C. H. Hollweg, and D. Lederman.

The fourth chapter considers the impact of short-run demand shocks emanating from major global markets on developing countries through the trade channel. A version of this chapter was published as a World Bank Policy Research Working Paper.

## Statements of Contributions

# Statement of Authorship

| | |
|---|---|
| Title of Paper | Structural Reforms and Labor Market Outcomes: International Panel-Data Evidence |
| Publication Status | ○ Published, ○ Accepted for Publication, ☒ Submitted for Publication, ○ Publication style |
| Publication Details | Submitted to the World Economy in June 2013 as: Hollweg, C. H, D. Lederman, and D. Mitra (2013). "Structural Reforms and Labor Market Outcomes: International Panel-Data Evidence." |

### Author Contributions

By signing the Statement of Authorship, each author certifies that their stated contribution to the publication is accurate and that permission is granted for the publication to be included in the candidate's thesis.

| | | | |
|---|---|---|---|
| Name of Principal Author (Candidate) | Claire H. Hollweg | | |
| Contribution to the Paper | Contributed to the planning of the article and the methodology, undertook literature review, collected data, performed analysis, helped in interpretation of the results, wrote parts of the manuscript, acted as corresponding author | | |
| Signature | | Date | Oct. 8, 2013 |

| | | | |
|---|---|---|---|
| Name of Co-Author | Daniel Lederman | | |
| Contribution to the Paper | contributed to the planning of the article and the methodology, supervised in development of the work, helped in interpretation of the results, provided comments on manuscript | | |
| Signature | | Date | 10/10/13 |

| | | | |
|---|---|---|---|
| Name of Co-Author | Devashish Mitra | | |
| Contribution to the Paper | contributed to the planning of the article and the methodology, supervised in development of the work, helped in interpretation of the results, wrote parts of the manuscript | | |
| Signature | | Date | 10/17/13 |

| | | | |
|---|---|---|---|
| Name of Co-Author | | | |
| Contribution to the Paper | | | |
| Signature | | Date | |

# Statement of Authorship

| | |
|---|---|
| Title of Paper | Measuring Capital Matters |
| Publication Status | O Published, O Accepted for Publication, O Submitted for Publication, X Publication style |
| Publication Details | Unpublished |

## Author Contributions

By signing the Statement of Authorship, each author certifies that their stated contribution to the publication is accurate and that permission is granted for the publication to be included in the candidate's thesis.

| | | | |
|---|---|---|---|
| Name of Principal Author (Candidate) | Claire H. Hollweg | | |
| Contribution to the Paper | Contributed to the planning of the article and the methodology, undertook literature review, collected data, performed analysis, helped in interpretation of the results, wrote manuscript | | |
| Signature | | Date | Oct. 8, 2013 |

| | | | |
|---|---|---|---|
| Name of Co-Author | Daniel Lederman | | |
| Contribution to the Paper | contributed to the planning of the article and the methodology, supervised in development of the work, helped in interpretation of the results | | |
| Signature | | Date | 10/10/13 |

| | | | |
|---|---|---|---|
| Name of Co-Author | | | |
| Contribution to the Paper | | | |
| Signature | | Date | |

| | | | |
|---|---|---|---|
| Name of Co-Author | | | |
| Contribution to the Paper | | | |
| Signature | | Date | |

# Statement of Authorship

| | |
|---|---|
| Title of Paper | Monitoring Export Vulnerability to Changes in Growth Rates of Major Global Markets |
| Publication Status | ☒ Published, ○ Accepted for Publication, ○ Submitted for Publication, ○ Publication style |
| Publication Details | Hollweg, C. H., D. Lederman, and J.-D. Reyes (2012). "Monitoring Export Vulnerability to Changes in Growth Rates of Major Global Markets." Policy Research Working Paper No. 6266, The World Bank, Washington, DC. |

## Author Contributions

By signing the Statement of Authorship, each author certifies that their stated contribution to the publication is accurate and that permission is granted for the publication to be included in the candidate's thesis.

| | | | |
|---|---|---|---|
| Name of Principal Author (Candidate) | Claire H. Hollweg | | |
| Contribution to the Paper | Contributed to the planning of the article and the methodology, undertook literature review, collected data, performed analysis, helped in interpretation of results, extended results and revised manuscript, wrote parts of the manuscript | | |
| Signature | | Date | Oct. 8, 2013 |

| | | | |
|---|---|---|---|
| Name of Co-Author | Daniel Lederman | | |
| Contribution to the Paper | Contributed to the planning of the article and the methodology, supervised in development of the work, helped in interpretation of the results, provided comments on manuscript | | |
| Signature | | Date | 10/10/13 |

| | | | |
|---|---|---|---|
| Name of Co-Author | José-Daniel Reyes | | |
| Contribution to the Paper | Contributed to the planning of the article and the methodology, collected data, performed parts of the analysis, helped in interpretation of the results, wrote parts of the manuscript, acted as corresponding author | | |
| Signature | | Date | Oct 21st, 2013 |

| | | | |
|---|---|---|---|
| Name of Co-Author | | | |
| Contribution to the Paper | | | |
| Signature | | Date | |

# Chapter 1

# Measuring Consumption Smoothing in Australia: An empirical test of the permanent income hypothesis

Claire H. Hollweg

# 1. Introduction and Motivation

Economists have long been interested in the relationship between income and consumption. In particular is the extent to which unexpected changes in income translate into changes in consumption, that is, whether individuals are able to smooth consumption against income shocks. Consumption is a direct measure of individual and household well-being. While income streams exhibit fluctuations, knowledge of the extent of consumption smoothing is informative about the welfare effects of shifts in the income distribution.

This study examines the link between individual-specific changes in income and changes in consumption of households in Australia over the period 2001-2009 using the Household Income and Labor Dynamics of Australia (HILDA) dataset. In particular, it estimates the degree of transmission of permanent as well as transitory idiosyncratic income shocks to consumption following the methodology of Blundell et al.'s (2008) study for the United States over the period 1979-1992. The degree of these transmissions, called the "partial consumption-smoothing parameters" of permanent and transitory shocks, are identified from the authors' permanent-transitory model that characterizes the processes of unexplained income growth and consumption growth. The model's parameters are then estimated using Australian household-level panel data. This analysis is disaggregated according to education, cohort of birth, and wealth to examine whether heterogeneity exists in the degree of consumption smoothing across different population subgroups. In addition, it empirically analyzes the mechanisms behind the degree of consumption smoothing found in the data, in particular the role of durable purchases, female labor supply, and taxes and transfers.

One of the unique characteristics of the HILDA is the household-level panel data on both consumption and income (as well as other information about economic and subjective well-being, labor market dynamics, and family dynamics). In fact, no such dataset is available for the United States. Rather, Blundell et al. (2008) imputed nondurable consumption data for households using data on food consumption and other household characteristics. Thus applying the HILDA survey data to Blundell et al.'s (2008) methodology is a key contribution of this study as it can estimate the parameters of interest with less error.

Blundell et al.'s (2008) model rests two extreme characterizations of individual behavior (as well as intermediate cases). At one extreme is the permanent income hypothesis in which personal savings is the only mechanism available to households to smooth consumption against idiosyncratic shocks to income. It predicts that changes in consumption are determined by permanent changes in income rather than transitory changes in income. Thus the nature of the relationship between income changes and consumption changes will depend on the degree of persistence of income shocks: transitory changes in income should have little effect on consumption, whereas permanent changes should have a significant effect. Thus if idiosyncratic shocks are persistent, no consumption smoothing scheme is necessary. At the other extreme is the complete markets hypothesis where individuals have full access to credit markets to smooth consumption. The model assumes that consumption is fully smoothed against transitory and permanent idiosyncratic shocks to income. When shocks are specific to the individual, then one way individuals can smooth consumption is via financial markets. (Other ways also exist, for example, through savings or durable goods purchases.) However, if shocks are macroeconomic and perfectly correlated among individuals, then individuals cannot write contracts to smooth consumption amongst themselves.[1]

For the entire sample of Australian households, nearly full consumption smoothing exists against transitory shocks that are specific to the individual. Although less consumption smoothing exists against permanent shocks, Australian households still achieve a high degree of consumption smoothing against highly persistent shocks, particularly when compared to households in the United States. The study's empirical results suggest that a 10 percent

[1] Alternatively, one may think of the rational-expectations permanent income hypothesis (where income follows a random walk if households have rational expectations and consume from permanent income) against the myoptic rule-of-thumb consumer (where consumption is not smoothed but instead follows current labor income).

permanent income shock induces about a 1 percent change in consumption. This suggests that for the average Australian household during the sample period, the marginal propensity to consume out of a permanent AUD 1 increase in income is about 30 cents. Although Blundell et al.'s (2008) estimate of the consumption-smoothing parameter of transitory shocks is similar in magnitude to that found in Australia, the estimate of the consumption-smoothing parameter of permanent shocks is significantly higher. In the United States, a 10 percent permanent income shock was found to induce about a 6 percent change in consumption. In addition, there is reason to believe Blundell et al.'s (2008) estimates for the United States are biased downwards due to measurement error in imputed consumption. The baseline results for Australia accord well with the complete markets hypothesis, which predicts full consumption smoothing to idiosyncratic income shocks.

When the analysis is disaggregated across different population subgroups, there is some support for consumption smoothing through precautionary savings. Permanent shocks are smoothed to a greater extent by older cohorts than younger cohorts (although younger cohorts are still able to partially smooth consumption against permanent shocks). Similar to the United States, low-wealth households in Australia are less able to smooth consumption against permanent shocks. Durable purchases, female labor supply, and taxes and transfers are all found to act as consumption-smoothing mechanisms.

Relatively few empirical tests of the permanent income hypothesis exist for Australia, as one would expect consumers to act similar in all countries in the rich world. In addition, all use aggregate consumption and income data with one noted exception. Earlier work in Australia tested the permanent income hypothesis's prediction that consumption should not respond to transitory changes in income. MacDonald and Kearney (1990) easily reject the permanent income hypothesis, while McKibbin and Richards (1988) and Johnson (1983) find evidence in favor of the hypothesis. Holloway (1992) tests the random walk component of the hypothesis:

if households have rational expectations and consume from permanent income, then consumption should follow and random walk. The author finds support for the hypothesis in that Australian data on aggregate consumption is well characterized as following a random walk (although with drift).

More closely related to this study are the works of Tan and Voss (2003) and Dvornak and Kohler (2007), which examine the relationship between consumption and wealth in Australia. Using a panel of Australian states for 1984-2001, Dvornak and Kohler (2007) find that changes in housing and stock market wealth have a significant effect on consumption expenditure: a permanent AUD 1 increase in stock market wealth increases annual consumption by about 9 cents, and the same increase in housing wealth increases annual consumption by about 3 cents. Tan and Voss (2003) have estimated that during the period 1988-1999, annual consumption increased by 4 cents in response to an AUD 1 increase in wealth. Using household-level data, Berger-Thomson et al. (2010) estimate the marginal propensity to consume in Australia due to two policy changes: income tax rates and lump-sum transfers (the baby bonus). While the marginal propensity to consume is found to be unity for tax cuts, it is 0.1 for lump-sum transfers.

The joint evolution of consumption and income inequality in Australia during this time period is also examined. The model's structure allows for the variances of permanent and transitory shocks to be estimated, as well as the extent of serial correlation of transitory shocks. It is therefore possible to examine whether changes in the persistence of income shocks have affected the joint evolution of income and consumption inequality. Consumption inequality is important if one is concerned with inequality in well-being rather than inequality in wages, earnings, or income, given that consumption provides a better measure of well-being than income. For example, consumption better reflects long-run resources while income measures fail to capture differences in income that result from differences in the accumulation of assets or access to credit (Meyer and Sullivan 2013). Understanding the link between income inequality and consumption inequality is also important for government policies, for example, such as the design of progressive income taxation.

Consumption inequality (measured as the variance of log household nondurable consumption) of Australian households increased slightly in the later part of the sample period. The estimation results show that the variance of the transitory shock changed little between 2002 and 2009, while the variance of the permanent shock increased nearly 50 percent between the beginning and end of the sample period. The results also show that households have less consumption smoothing against permanent versus transitory shocks. Together these results provide one explanation for the observed increase in consumption inequality: idiosyncratic income shocks were becoming more persistent, which households are less able to smooth consumption against.

The remainder of the chapter proceeds as follows. Section 2 discusses the principal results of the Blundell et al. (2008) study for the United States, as well as one methodological issue that this study is able to overcome. Section 3 presents the model and the procedure to estimate the consumption-smoothing parameters as well as the variances of the permanent and transitory income shocks, which are informative of changes in the persistence of income shocks. Section 4 summarizes the Australian household survey data and Section 5 the results. Section 6 concludes.

## 2. Discussion of Blundell et al. (2008)

Blundell et al. (2008) examine the joint evolution of income and consumption inequality of American households between 1979 and 1992, a period that witnessed large increases in income inequality in the United States. Measured as the variance of the log of income, income inequality grew by more than 30 percent in their sample of households between 1980 and

1992. But the authors also observe a divergence in the evolution of consumption and income inequality: while consumption and income inequality increased at the same rate from 1980 until 1985, consumption inequality stopped growing while income inequality continued to rise until 1988 before flattening off. The underlying question that the authors ask is whether changes in the income process or the nature of consumption smoothing may have driven this wedge between consumption and income inequality. In fact, their study is one of the first to try and analyze the link between changes in income inequality and changes in consumption inequality of American households during this period.[2]

The authors construct a model that allows for both permanent and transitory idiosyncratic shocks to the income-generating process, as well as partial smoothing of consumption to each of these income shocks (referred to in the paper as "partial insurance"). Assuming structures for the processes of permanent and transitory shocks (and thus income), and using panel data on income and (imputed) consumption, the variances of permanent and transitory income shocks as well as the degree of transmission of these shocks to consumption (the "partial consumption-smoothing parameters") are estimated. Furthermore, the variances of the shocks are allowed to differ by year and the estimates of the consumption-smoothing parameters are allowed to differ between the earlier and later part of the 1980s.[3] The evolution of the variances of permanent and transitory income shocks as well as the partial smoothing of consumption to each of these shocks are subsequently used to explain the change in the joint evolution of consumption and income inequality.

The authors conclude that the observed divergence between income and consumption inequality in the United States during the 1980s was due to a change in the persistence of income shocks coupled with asymmetric ability for households to smooth consumption against income shocks of different persistence. Growth in the variance of permanent shocks

[2] See Blundell et al. (2008) for an exhaustive review of the related literature.
[3] The model and estimation strategy is reviewed formally in Section 3 below.

was observed during the beginning of the 1980s, for which households were able to only partially smooth consumption against. In the late 1980s there is instead a growth in the variance of transitory shocks, for which households are better able to smooth consumption against. The authors did not find differences in the consumption-smoothing parameters in the earlier and later part of the sample. Thus it was not greater consumption smoothing opportunities that drove this wedge between income and consumption inequality, but growth in transitory income shocks that are more insurable.

The economic framework allows for consumption smoothing of transitory shocks through savings (the permanent income hypothesis), and consumption smoothing of all idiosyncratic shocks through complete markets (the complete markets hypothesis). As such, the estimates of the consumption-smoothing parameters can be used as a test of the validity of each of these models. However, for the United States, neither of these models was found to accord with the evidence. Households were found to fully smooth consumption against transitory shocks and partially smooth consumption against permanent shocks.

The degree of consumption smoothing to permanent shocks, however, was found to differ by cohort and wealth, where older cohorts and wealthier households were better able to smooth consumption against permanent shocks. This result supports the intermediate case of consumption smoothing through precautionary savings. Here, accumulated wealth is used to smooth consumption against persistent income shocks, thus high-income households and older cohorts who have accumulated more financial wealth are expected to have more consumption smoothing.

It has been argued that Blundell et al.'s (2008) estimates of the consumption-smoothing parameters should become the yardstick of quantitative macroeconomics (Kaplan and Violante 2010). For example, to measure whether macroeconomic models used for quantitative analysis admit the right amount of household consumption smoothing. However, there is one reason in particular to question the validity of Blundell et al.'s (2008) results. The empirical assessment of the transmission of income shocks into consumption requires panel data on income and on a comprehensive measure of consumption. In the United States, however, no such dataset is available. The Panel Study of Income Dynamics (PSID) contains information on income and food consumption, whereas the Consumer Expenditure Survey (CEX) contains detailed information on consumption but is not a panel dataset (rather it is a repeated cross-sectional dataset). To overcome this constraint, Blundell et al. (2008) construct a panel dataset for the United States using both the PSID and CEX that contains information on household income and imputed nondurable consumption.[4]

The authors exploit the fact that food consumption is available in both the PSID and the CEX. From the CEX, the authors estimate a demand function for food based on nondurable consumption expenditures, also allowing demand to change with relative prices and a host of demographic and socioeconomic characteristics of the household. Inverting the demand function then allows nondurable consumption to be imputed into the PSID for each household and year observation based on their food consumption reported in the PSID.

It is expected that this imputation procedure will lead to a downward biased estimate of the degree of consumption smoothing to income shocks (or an upward biased parameter estimate). The imputation measures consumption with error, and those measurement errors inflate the sample variance of the imputed consumption and subsequently the level of consumption inequality. Consumption appears more responsive in the data to changes in income than what actually occurs, thus underestimating the actual degree of consumption smoothing. In fact, Casado (2011) shows that the imputation procedure does indeed overestimate the true consumption response to permanent shocks.

[4] Other approaches in the literature have been to use only food consumption (for example, Hall and Mishkin (1982) and Hayashi et al. (1996)), or a more comprehensive consumption measure that lacks a panel dimension (for example, Blundell and Preston (1998)).

Thus a key contribution of this study is utilization of the HILDA dataset that contains data on both consumption and income. By reducing error in the consumption measure, this study provides more reliable estimates of the consumption-smoothing parameters.

## 3. Methodology

The primary purpose of the study's analysis is to identify the degree of transmission of permanent versus transitory income shocks to consumption, what Blundell et al. (2008) refer to as "partial insurance parameters" and what this study renames "partial consumption-smoothing parameters." To this end, the authors develop a permanent-transitory model to characterize the evolution of the unexplained income and consumption processes (assuming that the sole relevant source of idiosyncratic uncertainty faced by the consumer is net household income). The model can then be estimated using structural empirical techniques as to identify the consumption-smoothing parameters.

The model also allows for changes in the persistence of shocks to income by allowing the variances of the permanent and transitory income shocks to vary over time. These variances can be estimated as well as the extent of serial correlation of transitory shocks. This representation thus also provides a secondary goal, which is to understand the evolution in the consumption and income distributions and whether changes in the persistence of shocks affect the evolution of income and consumption inequality.

This section proceeds by first presenting the permanent-transitory model of Blundell et al. (2008). It next derives the moment restrictions that are used to identify the parameters of interest in the structural estimation strategy, which is described before proceeding to a critical discussion of the model.

## a. Model

The formalization of the model begins by placing structure of the income process of the household. Real (log) income of household $i$ at time $t$, $\log Y_{i,t}$, depends on a set of observable characteristics known to the consumer (such as education), $Z_{i,t}$, a permanent component, $P_{i,t}$, and a transitory component, $v_{i,t}$:

$$\log Y_{i,t} = Z'_{i,t}\phi_i + P_{i,t} + v_{i,t}.$$

Assume that the permanent component follows a martingale process of the form:

$$P_{i,t} = P_{i,t-1} + \zeta_{i,t}$$

where $\zeta_{i,t}$ is serially uncorrelated. Assume that the transitory component, $v_{i,t}$, follows an MA($q$) process of the form:

$$v_{i,t} = \sum_{j=0}^{q} \theta_j \epsilon_{i,t-j}$$

where $\theta_0 = 1$. Then (unexplained) income growth, $\Delta y_{i,t}$, is:

$$\Delta y_{i,t} = \zeta_{i,t} + \Delta v_{i,t} \tag{1}$$

where $y_{i,t} = \log Y_{i,t} - Z'_{i,t}\phi_i$ denotes the log of real income net of predictable individual components.

To consider the degree of transmission of income shocks to consumption, the model next defines the (unexplained) change in log consumption, $\Delta c_{i,t}$, as:

$$\Delta c_{i,t} = \phi_i \zeta_{i,t} + \psi_i \epsilon_{i,t} + \xi_{i,t} \tag{2}$$

where $c_{i,t} = \log C_{i,t} - Z'_{i,t}\phi_i$ is the log of real consumption net of its predictable components and the random term $\xi_{i,t}$ represents innovations in consumption that are independent of those to income, for example, preference shocks. The parameters $\phi_i$ and $\psi_i$ are the consumption-smoothing parameters, as they reflect the degree of transmission of permanent income shocks $\zeta_{i,t}$ and transitory income shocks $\varepsilon_{i,t}$ to consumption, respectively.

Thus Equation (2) nests two extreme cases of consumption smoothing: full consumption smoothing ($\phi_i = \psi_i = 0$) and no consumption smoothing ($\phi_i = \psi_i = 1$). Intermediate cases can also arise in which $0 < \phi_i < 1$ and $0 < \psi_i < 1$. A higher degree of consumption smoothing exists the closer the coefficients are to zero. In addition, the model allows the degree of transmission to vary across individuals. For example, allowing consumption-smoothing parameters to differ by birth cohort allows differences to be interpreted as age effects.

Assuming $\xi_{i,t}$ is stationary, one can write the following decomposition for the change in the variance of consumption growth:

$$\Delta\text{var}(\Delta c_t) = \text{var}(\zeta_t)\Delta\phi^2 + \phi^2\Delta\text{var}(\zeta_t) + \text{var}(\epsilon_t)\Delta\psi^2 + \psi^2\Delta\text{var}(\epsilon_t). \qquad (3)$$

Equation (3) shows that the variance of consumption growth can increase due to a decline in the degree of consumption smoothing with respect to income shocks (for given variances), implying the values for $\phi$ or $\psi$ have increased, or because of an increase in the variances of income shocks (for given consumption smoothing). Ideally the analysis would be able to separate the different forces at play visible in this equation. However, there are not enough time periods in the sample to allow the consumption-smoothing parameters to vary over time. Assuming these are constant during the sample period, then, shows that the variance of consumption growth can increase because the variance of the transitory shock or permanent shock (or both) increased, conditional on households not having full consumption smoothing against these shocks. This decomposition is important for the interpretation of the results below.

## b. Deriving covariance restrictions

Covariance restrictions imposed on the unexplained income and consumption growth processes, $\Delta y_{i,t}$ and $\Delta c_{i,t}$, given by equations (1) and (2), respectively, are used to estimate the parameters of interest when panel data is available. These moments can be computed for the whole sample or for individuals belonging to a homogenous group (for example, by cohort or education level).

In addition, the framework allows consumption to be measured with error. Respondents are asked to keep journals of consumption expenditure, and it is highly likely that these records will not be measured with the level of accuracy in which income is measured. However, this study still argues that any measurement error will be less than what would be present with imputed consumption. Let $c_{i,t}^*$ be measured consumption and $c_{i,t}$ be true consumption. Then:

$$c_{i,t}^* = c_{i,t} + u_{i,t}$$

where $u_{i,t}$ is measurement error. Even under the assumption that consumption follows a martingale process, this measurement error will create serial correlation in observed consumption growth.

Let $\mathrm{var}(\cdot)$ and $\mathrm{cov}(\cdot,\cdot)$ denote cross-sectional variances and covariances, respectively. Assume that $\zeta_{i,t}$, $v_{i,t}$, and $\xi_{i,t}$ are mutually uncorrelated processes. Then the following covariance restrictions of unexplained income growth can be derived from Equation (1) for different lags:

$$\mathrm{cov}(\Delta y_t, \Delta y_{t+s}) = \begin{cases} \mathrm{var}(\zeta_t) + \mathrm{var}(\Delta v_t) & \text{for } s = 0 \\ \mathrm{cov}(\Delta v_t, \Delta v_{t+s}) & \text{for } s \neq 0 \end{cases}$$

where $\mathrm{cov}(\Delta y_t, \Delta y_{t+s})$ and $\mathrm{var}(\Delta y_t)$ are the autocovariance and variance of income growth, respectively. This equation shows that income inequality (obtained by setting $s = 0$) may increase either because of increases in the variance of permanent shocks, or because of an increase in the variance of income growth due to transitory shocks. In addition, higher order autocovariances inform on the transitory and persistent nature of such shifts. This discussion is taken up in the results section below.

The covariance term $\text{cov}(\Delta v_t, \Delta v_{t+s})$ depends on the serial correlation properties of $v$. This study adopts an MA(1) process for the transitory component of income.[5] Under this assumption, the specific forms of the covariance restrictions of income growth can be derived as:

$$\text{cov}(\Delta y_t, \Delta y_{t+s}) = \begin{cases} \text{var}(\zeta_t) + var(\epsilon_t) + (1-\theta)^2\text{var}(\epsilon_{t-1}) + \theta^2\text{var}(\epsilon_{t-2}) & \text{for } s = 0 \\ -(1-\theta)\text{var}(\epsilon_t) + (1-\theta)\theta\text{var}(\epsilon_{t-1}) & \text{for } |s| = 1 \\ -\theta\text{var}(\epsilon_t) & \text{for } |s| = 2 \\ 0 & \text{for } |s| > 2 \end{cases}$$

The same approach is taken for consumption where Equation (2) can be used to derive the covariance restrictions of unexplained consumption growth for different lags:

$$\text{cov}(\Delta c_t^*, \Delta c_{t+s}^*) = \begin{cases} \phi^2\text{var}(\zeta_t) + \psi^2\text{var}(\epsilon_t) + \text{var}(\xi_t) + \text{var}(u_t) + \text{var}(u_{t-1}) & \text{for } s = 0 \\ -\text{var}(u_t) & \text{for } s \neq 0 \end{cases}$$

where $\text{var}(u_t)$ is the variance of measurement error. Due to the consumption martingale assumption, the variance of the measurement error can be identified as the first order autocovariance of unexplained consumption growth (since any serial correlation in consumption growth should only be attributed to noise). As pointed out above, this equation shows that consumption inequality ($s = 0$) can increase due to a decline in the degree of insurance with respect to income shocks (for given variances) or an increase in the variances of income shocks (for given insurance) (or an increase in the variance of consumption measurement error).

Equations (1) and (2) together can be used to derive the covariance restrictions between unexpected income growth and unexpected consumption growth at various lags:

$$\text{cov}(\Delta c_t^*, \Delta y_{t+s}) = \begin{cases} \phi\text{var}(\zeta_t) + \psi\text{var}(\epsilon_t) & \text{for } s = 0 \\ \psi\text{cov}(\epsilon_t, \Delta v_{t+s}) & \text{for } s > 0 \end{cases}$$

where $\text{cov}(\Delta c_t^*, \Delta y_{t+s})$ is the joint income and consumption moment of unexplained income and consumption growth (allowing for measurement error in consumption). This moment

[5] Identification of the serial correlation coefficients $\theta_j$ does not hinge on the order of the process $q$. Allowing for an MA($q$) process, for example, adds $q-1$ extra parameters (the $q-1$ MA coefficients) but also $q-1$ extra moments, so that identification is unaffected.

reflects the extent of consumption smoothing with respect to transitory shocks, and would equal zero if the household had full consumption smoothing of transitory shocks.

Under the assumption of an MA(1) process for the transitory component of income, the specific forms of the covariance restrictions of unexplained income and consumption growth can be derived as:

$$\operatorname{cov}(\Delta c_t^*, \Delta y_{t+s}) = \begin{cases} \phi \operatorname{var}(\zeta_t) + \psi \operatorname{var}(\epsilon_t) & \text{for } s = 0 \\ -\psi(1-\theta)\operatorname{var}(\epsilon_t) & \text{for } s = 1 \\ -\psi\theta \operatorname{var}(\epsilon_t) & \text{for } s = 2 \\ 0 & \text{for } s > 2 \text{ and } s < 0 \end{cases}.$$

## c. Estimation method

Panel data is used to estimate the parameters of the model, which are: variances of permanent ($\zeta$) and transitory ($\epsilon$) income shocks, the extent of serial correlation in transitory shocks ($\theta$), the consumption-smoothing parameters ($\phi$ and $\psi$), the variance of unobserved heterogeneity ($\xi$), and measurement error ($u$). This is done for the entire sample as well as for different subgroups of the population.

The availability of panel data presents several advantages over a repeated cross-sectional analysis, as discussed in Blundell et al. (2008). First, with repeated cross sections, the variances and covariances of income and consumption growth cannot be observed. Second, with panel data, one can estimate a richer model with the consumption-smoothing parameters left free and thus test the validity of alternative explanations regarding the evolution of consumption inequality over time. Third, although it is not necessary to have panel data on both consumption and income to identify the variance of the income shocks, using panel data on both consumption and income improves efficiency of these estimates because it provides additional moments for identification.

The structural estimation strategy uses diagonally weighted minimum distance (DWMD) estimation. DWMD is a simple generalization of equally weighted minimum distance but allows for heteroskedasticity. In a general sense, the parameters of the model are chosen as to minimize the distance between the actual moments in the joint panel data on income and consumption growth and the theoretical moment restrictions derived from the variance-covariance structure of income and consumption growth (given above in section 3.2). The estimation method is undertaken as follows.

For each individual $i$, define two vectors of interest as:

$$\Delta c_i^* = \begin{pmatrix} \Delta c_{i,1}^* \\ \Delta c_{i,2}^* \\ \cdots \\ \Delta c_{i,T}^* \end{pmatrix} \qquad \Delta y_i = \begin{pmatrix} \Delta y_{i,1} \\ \Delta y_{i,2} \\ \cdots \\ \Delta y_{i,T} \end{pmatrix}$$

where $\Delta c_{i,t}^*$ and $\Delta y_{i,t}$ are unexplained (measured) consumption growth and income growth between years $t-1$ and $t$, respectively. These are obtained as the differences in the residuals of the regression of real log consumption and log income on the set of observable characteristics $Z_{i,t}$. Let $t = 0$ indicate the first year of the household-level consumption and income data and $t = T$ indicate the last year. Then $\Delta c_i^*$ and $\Delta y_i$ are of dimension $T$. Moreover, if the individual was not observed in year $t$ or $t-1$, the unobservable $\Delta c_{i,t}^*$ and $\Delta y_{i,t}$ are replaced with zeros.[6]

With these vectors define:

$$d_i^{c^*} = \begin{pmatrix} d_{i,1}^{c^*} \\ d_{i,2}^{c^*} \\ \cdots \\ d_{i,T}^{c^*} \end{pmatrix} \qquad d_i^y = \begin{pmatrix} d_{i,1}^y \\ d_{i,2}^y \\ \cdots \\ d_{i,T}^y \end{pmatrix}$$

[6] Since the Australian household panel dataset has missing years for which consumption data was collected (the number of years depending on the definition of consumption), the vector $\Delta c_i^*$ is understood to be dimension $T - t_m$ where $t_m$ is the number of missing years. That is, the rows with missing consumption data have already been swept out from $\Delta c_i^*$. For example, non-durable consumption is available for 2005-2009 thus $\Delta c_i^*$ is of diminution 4.

where $d_{i,t}^{c^*} = 1$ if $\Delta c_{i,t}^*$ is not missing and similarly $d_{i,t}^{y} = 1$ if $\Delta y_{i,t}$ is not missing. Then $d_i^{c^*}$ and $d_i^{y}$ are also of dimension $T$. This notation allows in a simple manner to handle the problems of unbalanced panel data.

Stacking observations in $\Delta c^*$ and $\Delta y$ and in $d^{c^*}$ and $d^{y}$ for each individual yields the vectors:

$$x_i = \begin{pmatrix} \Delta c_i^* \\ \Delta y_i \end{pmatrix} \qquad d_i = \begin{pmatrix} d_i^{c^*} \\ d_i^{y} \end{pmatrix}$$

which are each of dimension $2T$.

With these vectors it is possible to derive the vector of moments as:

$$m = \text{vech}\left\{\left(\sum_{i=1}^{N} x_i x_i'\right) \oslash \left(\sum_{i=1}^{N} d_i d_i'\right)\right\}$$

which is of dimension $T^2$ and contains $T(2T+1)$ moments.[7] This vector $m$ contains the estimates of the covariances of interest identified above, including $\text{cov}(\Delta y_t, \Delta y_{t+s})$ , $\text{cov}(\Delta c_t^*, \Delta c_{t+s}^*)$, and $\text{cov}(\Delta c_t^*, \Delta y_{t+s})$.

The variance-covariance matrix of $m$ is

$$V = \left(\sum_{i=1}^{N} \left((m_i - m)(m_i - m)'\right) \otimes (D_i D_i')\right) \oslash \left(\sum_{i=1}^{N} D_i D_i'\right)$$

where $D_i = \text{vech}\{d_i d_i'\}$ and $m_i = \text{vech}\{x_i x_i'\}$. The square roots of the elements in the main diagonal of $V$ provide the standard errors of the corresponding elements in $m$.

The estimated model for $m$ is given as:

$$m = f(\Lambda)$$

where $\Lambda$ is the vector of parameters of interest. This model is defined by the covariance restrictions derived above that are functions of the underlying parameters of interest. Specifically, the parameters of interest in $\Lambda$ are estimated by minimizing

[7] In practice there are fewer than $T(2T+1)$ moments because data on consumption is not available in all years.

$$\min_{\Lambda}\big(m - f(\Lambda)\big)' A(m - f(\Lambda))$$

where the weighting matrix $A$ is a diagonal matrix with $\text{diag}(V^{-1})$.[8]

For inference purposes it is necessary to compute of standard errors, given by

$$\widehat{\text{var}(\hat{\Lambda})} = (G'AG)^{-1}G'AVAG(G'AG)^{-1}$$

where $G = \partial f(\Lambda)/\partial\Lambda|_{\Lambda=\hat{\Lambda}}$ is the Jacobian matrix evaluated at the estimated parameters $\hat{\Lambda}$.

## d. Critical discussion

In recent years there has been a resurgence of interest in estimating the degree of consumption smoothing, either using quasi-experimental data or imposing structural restrictions on the stochastic income process faced by consumers. The study's estimation strategy focuses exclusively on moments based on income and consumption growth to identify the extent of consumption smoothing to income shocks. This is a standard approach adopted in the literature.[9] However, structural estimation techniques are not without criticism. The structural approach specifies a complete model of economic behavior then estimates or calibrates the parameters of the model. Critics of the structural approach argue that it is difficult to identify all parameters in an empirically compelling manner due to, for example, selection effects, simultaneity, or omitted variables.

Instead, reduced form strategies that estimate statistical relationships can perform better in identification of causal effects. What makes this difficult, however, is that individual shocks cannot be directly observed in income data. Income changes are best described by a combination of highly persistent and highly transitory shocks (McCurdy 1982, Abowd and Card 1989, Blundell and Preston 1998). However, one observes only the total income change

[8] In practice, I write a function in Matlab that defines the $T$-by-$T$ matrix of covariance restrictions derived above, which is dependent on $f(\Lambda)$. I then minimize the distance between my actual vector of moments $m$ calculated from the data and the model $f(\Lambda)$ with respect to the argument $\Lambda$, which is the vector of parameters of interest.

[9] See, for example, Hall and Mishkin (1982), Blundell and Preston (1998), Guvenen and Smith (2008), and Primiceri and van Rens (2009).

and cannot disentangle the realization of the shocks of different persistence. As a consequence, some authors have chosen to simply measure the response of consumption to total income changes, whereas others have used proxies for permanent and transitory income changes in an attempt to separately identify the two shocks (such as randomization of the timing when tax rebate checks are received by households, as studied by Souleles et al. (2006)). However, one objection of these quasi-experimental studies is that the results may be context-specific.

Nevertheless, Blundell et al.'s (2008) framework does make some important progress in overcoming this difficulty. What is useful about the authors' structural approach is that it decomposes innovations to household income into transitory and permanent components by modeling the data-generating process for income. And by modeling the data-generating process for consumption, it can consider how changes in the variance of permanent and transitory income components are translated into changes in the variance of consumption. In addition, the fraction of permanent and transitory income shocks that are effectively smoothed by the individual can be estimated. However, one must make assumptions about the particular specifications of these processes, which leaves the model open to further criticism. One important cause for bias in the estimates is when the models themselves are misspecified, for example, the structure of the data-generating processes.

In the context of this study, the chosen income process follows a random walk for the permanent component and an MA(1) process for the transitory component. If, for example, the true structure of the income generating process was not characterized by a permanent component, but instead was characterized by a transitory component with high persistence, this would place a bias on the estimate of the degree of consumption smoothing.

The same is true for the consumption-generating process. The structural model maintains that consumption is martingale, which implies that consumption growth is orthogonal to past and future income shocks. Thus households have no foresight of future shocks and no history dependence of the consumption allocation with respect to past shocks. The consumption-smoothing parameter estimates will be biased if this assumption is not satisfied. One reason these orthogonality conditions may fail is if agents are liquidity constrained (Kaplan and Violante 2010). For example, if a household receives a negative transitory shock, such a household would like to borrow to smooth the negative shock, but cannot. Thus consumption is expected to drop the same period the agent receives the negative shock, but eventually will recover to its baseline level. And because agents prefer smooth paths for consumption, this adjustment takes place gradually. This therefore leads to a positive expected change in consumption over future periods.

## 4. Data

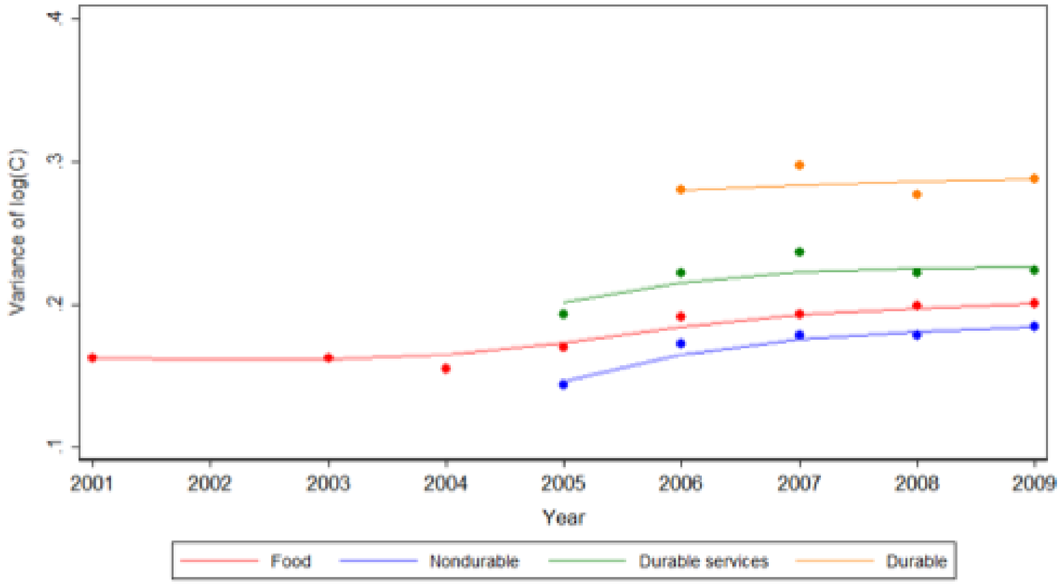
Data on Australian household-level consumption and income are collected from the HILDA annual survey for the sample period 2001 to 2009 (waves 1 to 9). The HILDA survey is a household-based panel study that began in 2001 with the first wave consisting of 7,682 households and 19,914 individuals. The individuals are followed over time and interviews are conducted annually with all adult members of each household.

One of the unique characteristics of the HILDA is the household-level panel data on both consumption and income (as well as other information about economic and subjective well-being, labor market dynamics, and family dynamics). In fact, no such dataset is available for the United States. As discussed above, Blundell et al. (2008) imputed nondurable consumption data for households in the PSID for which food consumption data was available after estimating demand equations from different cross-sections of the CEX. Thus applying the HILDA survey data to Blundell et al.'s (2008) methodology is a key contribution of this study as it can estimate the parameters of interest without needing to impute consumption data for households. This increases the reliability of the estimates by reducing measurement error.

Following Blundell et al. (2008), four separate measures of consumption are used. Food consumption is defined as the sum of annual expenditure on food at home and food away from home. Nondurable consumption, which is treated as the baseline consumption measure, is defined as the sum of food, alcohol, tobacco, and expenditure on other nondurable goods. This includes services, heating fuel, public and private transport (including gasoline), personal care, and semidurables of clothing and footwear. This baseline definition excludes expenditure on various durables (housing, furniture, appliances, etc.), health, and education. While food consumption data is available for nearly the entire sample period (with the exception of wave 2 in 2002), nondurable consumption data is only available for 2005-2009 (wave 5 onwards). To test the sensitivity of the results to the consumption measure as well as examine the role of different consumption-smoothing mechanisms, the results were also obtained using two additional definitions. First is a definition of nondurable consumption that includes services from some durables (housing and vehicles). Second is a definition of consumption that includes durables. Durable consumption data is available for 2006 to 2009 (wave 6 onwards). The data is then deflated by the CPI to real 2002 Australian dollars.

Figure 1 plots the variance of the log of the four definitions of consumption, which is the measure of consumption inequality adopted in this study. The graph plots the actual estimates of the variances as well as smoothing curves passing through the scatter. The variance of food and nondurable consumption steadily increased between the first year of data collection and 2009, while the slope of the durable consumption measure remained flat. Inequality of durable consumption is higher than that of nondurable (both definitions) and food consumption, and nondurable consumption posts the lowest variance of the consumption measures.

**Figure 1: Variance of log(C) measures**

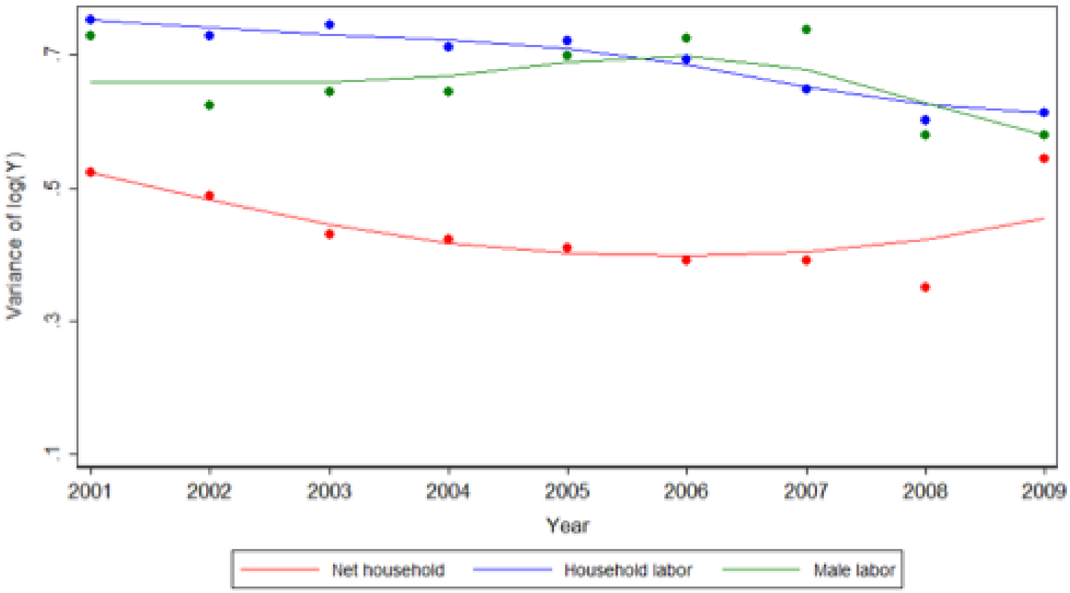



Notes: Figure 1 plots consumption inequality (the variance of the log) of four definitions of consumption (food, nondurable, durable services, durable), including the actual estimates of the variances in each year data is available and smoothing curves passing through the scatter.

The baseline income measure is net total household earnings, defined as household financial year gross wages and salaries plus public and private transfers such as welfare payments (but excluding income from financial assets) less federal taxes on non-financial income. Income is available for 2001 to 2009 (all waves of the HILDA survey). The results also experiment with two alternative definitions of income (following Blundell et al. (2008)), including total household labor earnings (excluding transfers and taxes) and male labor earnings. Adopting these alternative definitions helps examine the role of different consumption-smoothing mechanisms, such as taxes and transfers and female labor supply. All measures are deflated to real 2002 Australian dollars.

Figure 2 plots the variance of the log of the three definitions of income, which is the measure of income inequality. The graph plots the actual estimates of the variances as well as smoothing curves passing through the scatter. Not surprisingly, the variance of labor earnings is greater than that of net household earnings, highlighting the redistributive nature of taxes and transfers. However, trends in these variances are noticeably different. The variance of household net income and household labor income declined throughout the sample period, although at different rates, except for a large jump in 2009 in net household income. The variance of male labor income steadily increased between 2001 and 2007 then declined significantly in 2008 and 2009.

**Figure 2: Variance of log(Y) measures**




Notes: Figure 2 plots income inequality (the variance of the log) of three definitions of income (net household, household labor, male labor), including the actual estimates of the variances in each year data is available and smoothing curves passing through the scatter.

The sample consists of continuously married non-same sex couples with or without children headed by a male age 30 to 65. The household head is defined as the person who owns or rents the unit. Restricting the age of the household head avoids capturing consumption smoothing responses to changes in family composition and education if under 30 and retirement if over 65. In addition, households whose head or head's spouse changes are dropped from the sample. The sample selection is therefore chosen to focus on consumption-smoothing responses to income risk, not divorce, widowhood, or other household break-up factors (Blundell et al. (2008)).

Households were also dropped from the sample if there was missing information on: education, number of children, food expenditure, and incomplete consumption and income responses. Outliers were also dropped including households with zero before-tax income, zero total nondurable expenditure, income growth above 500 percent and below -80 percent, or with a level of income below AUD 100 in a given year.

It is questionable whether stable families have more or less access to consumption-smoothing mechanisms than unstable families, in which case sample selection bias could be an issue. On the one hand, stable families often have more income and assets and therefore are less likely to be eligible for social insurance, which is typically means-tested. On the other hand, they can plausibly be more successful in securing access to credit, family networks, and other informal consumption-smoothing devices, over and above consumption smoothing through savings.

The observable characteristics of the household include: dummies for birth cohort of the household head (equal to 1 if the household head was born in the 1940s, 1950s, 1960s, or 1970s); education as a dummy if household head completed a graduate degree (such as college or trade degree); number of household members; dummies for number of resident and non-resident children (1, 2, or more); dummy if household head worked in the year; dummy if household head's spouse worked in the year; dummy if there were other income recipients other than the household head and head's spouse; dummy if the household head is aboriginal or Torres Strait Islander; dummy if the household resides in a major urban area; and state of residence dummies.[10] Summary statistics are presented in Table 1.

**Table 1: Summary Statistics**

| | 2001 | 2002 | 2003 | 2004 | 2005 | 2006 | 2007 | 2008 | 2009 |
|---|---|---|---|---|---|---|---|---|---|
| Birth cohort 1940s | 0.233 | 0.232 | 0.227 | 0.224 | 0.197 | 0.174 | 0.152 | 0.135 | 0.124 |
| Birth cohort 1950s | 0.324 | 0.326 | 0.321 | 0.313 | 0.315 | 0.313 | 0.307 | 0.302 | 0.306 |
| Birth cohort 1960s | 0.326 | 0.330 | 0.329 | 0.339 | 0.339 | 0.335 | 0.337 | 0.332 | 0.320 |
| Birth cohort 1970s | 0.056 | 0.074 | 0.103 | 0.124 | 0.149 | 0.178 | 0.203 | 0.230 | 0.249 |
| College degree | 0.644 | 0.660 | 0.679 | 0.687 | 0.699 | 0.707 | 0.725 | 0.724 | 0.731 |
| Hh size | 3.591 | 3.581 | 3.569 | 3.564 | 3.550 | 3.569 | 3.564 | 3.551 | 3.503 |
| 1 child | 0.124 | 0.114 | 0.122 | 0.120 | 0.111 | 0.109 | 0.113 | 0.119 | 0.125 |
| 2 children | 0.391 | 0.405 | 0.407 | 0.414 | 0.411 | 0.414 | 0.415 | 0.396 | 0.411 |
| 3+ children | 0.418 | 0.410 | 0.399 | 0.399 | 0.399 | 0.390 | 0.389 | 0.399 | 0.382 |
| Hh head working | 0.779 | 0.783 | 0.805 | 0.801 | 0.811 | 0.816 | 0.827 | 0.835 | 0.820 |
| Hh head spouse working | 0.637 | 0.667 | 0.666 | 0.677 | 0.706 | 0.725 | 0.743 | 0.754 | 0.734 |
| Other income recipients | 0.183 | 0.190 | 0.201 | 0.211 | 0.231 | 0.236 | 0.243 | 0.263 | 0.258 |
| Aboriginal | 0.010 | 0.012 | 0.010 | 0.008 | 0.007 | 0.007 | 0.008 | 0.009 | 0.008 |
| Large city | 0.585 | 0.604 | 0.596 | 0.592 | 0.589 | 0.600 | 0.592 | 0.594 | 0.590 |

[10] Census collection districts are classified as major urban, other urban, bounded locality, rural balance, and migratory areas using population counts from the 2001 census.

| | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| Victoria | 0.260 | 0.254 | 0.249 | 0.254 | 0.248 | 0.251 | 0.248 | 0.251 | 0.249 |
| Queensland | 0.199 | 0.199 | 0.209 | 0.200 | 0.213 | 0.209 | 0.204 | 0.217 | 0.217 |
| South Aus | 0.094 | 0.096 | 0.091 | 0.098 | 0.091 | 0.095 | 0.094 | 0.090 | 0.099 |
| Western Aus | 0.097 | 0.095 | 0.095 | 0.101 | 0.096 | 0.098 | 0.092 | 0.093 | 0.090 |
| Tasmania | 0.031 | 0.028 | 0.029 | 0.031 | 0.036 | 0.033 | 0.035 | 0.032 | 0.028 |
| Northern Territory | 0.005 | 0.004 | 0.006 | 0.006 | 0.007 | 0.005 | 0.006 | 0.007 | 0.005 |
| Aus Capital Territory | 0.020 | 0.020 | 0.024 | 0.022 | 0.021 | 0.022 | 0.023 | 0.027 | 0.026 |
| New South Whales | 0.293 | 0.304 | 0.299 | 0.287 | 0.287 | 0.287 | 0.298 | 0.283 | 0.284 |
| Nondurable consumption | n.a. | n.a. | n.a. | n.a. | 19,819 | 23,415 | 23,627 | 24,444 | 23,630 |
| Food consumption | 10,730 | n.a. | 10,771 | 10,888 | 10,168 | 11,326 | 11,410 | 11,766 | 11,563 |
| Durable services consumption | n.a. | n.a. | n.a. | n.a. | 12,382 | 29,443 | 30,053 | 30,272 | 29,261 |
| Durable consumption | n.a. | n.a. | n.a. | n.a. | n.a. | 39,654 | 40,283 | 40,443 | 39,701 |
| Net hh income | 53,710 | 56,528 | 56,329 | 57,608 | 60,784 | 63,968 | 68,900 | 70,385 | 73,813 |
| Hh labor income | 62,432 | 66,083 | 66,935 | 67,982 | 72,041 | 75,731 | 80,655 | 81,502 | 82,217 |
| Male labor income | 43,495 | 45,037 | 46,219 | 46,811 | 48,506 | 50,003 | 53,670 | 53,416 | 53,795 |

Notes: Table 1 presents, in each year, the mean of each variable over included households from the HILDA survey.

# 5. Results

The discussion begins by presenting the regression results of real log income and consumption on the set of observable characteristics of the household. Following is an empirical characterization of the autocovariances of unexplained income and consumption growth (calculated as the year-on-year differences in the residuals of the log of income and consumption regressions). Finally, the DWMD estimates of the size and trends in the variance of permanent and transitory shocks to income and the degree of consumption smoothing to these shocks are discussed. Estimation results are first presented using the benchmark nondurable consumption and net household income measures. The analysis is then extended to explore how these results change for alternative definitions of consumption and income as well as for different subgroups of the population.

## a. Calculating unexplained consumption and income growth

Unexpected consumption (income) is the log of real consumption net of its predictable components. An ordinary least squares (OLS) regression of real log consumption (income) on the observable characteristics of the household, $Z_{i,t}$, removes the impact of the deterministic effects of the household on consumption choices (income outcomes). Unexpected consumption (income) growth is then calculated as the year-on-year difference in the residuals of the log of consumption (income) regression.

The household characteristics include: dummies for birth cohort of the household head; dummy for education of household head; family size; dummies for number of children; dummies for household head working, household head's spouse working, and other income recipients; dummy for race of household head; dummy if household is resident in a major urban area; and state of residence dummies. Year dummies are used to control for unobserved factors that affect consumption decisions and income outcomes of all households equally in a particular year. Interacting the variables for household head or head's spouse working, other income recipients, residence in a large city, and state allows the effects of these characteristics to vary over time. Tables 2a and 2b present the OLS regression results for the consumption and income measures, respectively. Coefficients on the interaction terms are omitted to save space.

## Table 2a: Consumption

| | Nondurable | Food | Durable services | Durables |
|---|---|---|---|---|
| Birth cohort 1950s | 0.007 | 0.032*** | -0.007 | 0.026 |
| | (0.0168) | (0.0112) | (0.0194) | (0.0264) |
| Birth cohort 1960s | 0.025 | 0.004 | 0.021 | 0.045 |
| | (0.0183) | (0.0124) | (0.0211) | (0.0288) |
| Birth cohort 1970s | -0.075*** | -0.078*** | -0.068*** | -0.055* |
| | (0.0206) | (0.0145) | (0.0236) | (0.0320) |
| College degree | 0.091*** | 0.058*** | 0.118*** | 0.149*** |
| | (0.0112) | (0.0077) | (0.0129) | (0.0173) |
| Hh size | 0.083*** | 0.117*** | 0.078*** | 0.081*** |
| | (0.0055) | (0.0039) | (0.0063) | (0.0085) |
| 1 child | -0.099*** | -0.058*** | -0.111*** | -0.078** |
| | (0.0227) | (0.0163) | (0.0261) | (0.0348) |
| 2 children | -0.082*** | -0.050*** | -0.080*** | -0.041 |
| | (0.0210) | (0.0151) | (0.0242) | (0.0325) |
| 3+ children | -0.121*** | -0.074*** | -0.145*** | -0.119*** |
| | (0.0234) | (0.0168) | (0.0269) | (0.0362) |
| Hh head working | 0.073*** | 0.097*** | 0.109*** | 0.149*** |
| | (0.0277) | (0.0266) | (0.0325) | (0.0374) |
| Hh head spouse working | 0.120*** | 0.085*** | 0.134*** | 0.169*** |
| | (0.0238) | (0.0228) | (0.0273) | (0.0320) |
| Other income recipients | 0.053** | 0.038* | 0.079*** | 0.035 |
| | (0.0240) | (0.0223) | (0.0275) | (0.0320) |
| Aboriginal | 0.066 | 0.019 | 0.029 | -0.062 |
| | (0.0557) | (0.0384) | (0.0650) | (0.0832) |
| Large city | 0.040* | 0.053*** | 0.035 | 0.041 |
| | (0.0214) | (0.0198) | (0.0246) | (0.0283) |
| Victoria | -0.004 | -0.047** | 0.006 | 0.021 |
| | (0.0251) | (0.0237) | (0.0288) | (0.0340) |
| Queensland | -0.048*** | -0.059*** | -0.031* | 0.006 |
| | (0.0142) | (0.0099) | (0.0163) | (0.0217) |
| South Aus | -0.073** | -0.099*** | -0.093** | -0.074 |
| | (0.0372) | (0.0335) | (0.0404) | (0.0468) |
| Western Aus | 0.089** | 0.112*** | 0.074* | 0.114** |
| | (0.0365) | (0.0342) | (0.0421) | (0.0479) |
| Tasmania | -0.080 | -0.098* | -0.178** | -0.056 |
| | (0.0585) | (0.0577) | (0.0709) | (0.0770) |
| Northern Territory | 0.194 | 0.117 | -0.021 | 0.092 |
| | (0.1622) | (0.1292) | (0.1610) | (0.1711) |
| Aus Capital Territory | -0.006 | -0.005 | -0.028 | -0.033 |
| | (0.0688) | (0.0634) | (0.0786) | (0.0847) |
| Constant | 9.544*** | 8.721*** | 9.692*** | 9.884*** |
| | (0.0372) | (0.0342) | (0.0450) | (0.0525) |
| Observations | 6,397 | 12,419 | 6,321 | 4,498 |
| R-squared | 0.140 | 0.199 | 0.114 | 0.116 |

Notes: Table 2a presents the estimated OLS regression coefficients for each of the consumption measures as the dependent variable in each column. Coefficients on the interaction terms are omitted to save space. Standard errors are in parentheses. *** significant at the 1% level; ** significant at the 5% level; * significant at the 10% level.

## Table 2b: Income

| | Net hh income | Hh labor income | Male labor income |
|---|---|---|---|
| Birth cohort 1950s | -0.030*** | 0.037** | 0.080*** |
| | (0.0114) | (0.0174) | (0.0212) |
| Birth cohort 1960s | -0.049*** | 0.022 | 0.112*** |
| | (0.0127) | (0.0194) | (0.0235) |
| Birth cohort 1970s | -0.038** | 0.011 | 0.142*** |
| | (0.0155) | (0.0233) | (0.0278) |
| College degree | 0.058*** | 0.158*** | 0.196*** |
| | (0.0080) | (0.0121) | (0.0144) |
| Hh size | 0.090*** | 0.006 | 0.010 |
| | (0.0040) | (0.0062) | (0.0074) |
| 1 child | -0.134*** | -0.087*** | 0.076** |
| | (0.0172) | (0.0254) | (0.0299) |
| 2 children | -0.186*** | -0.099*** | 0.100*** |
| | (0.0160) | (0.0238) | (0.0283) |
| 3+ children | -0.219*** | -0.175*** | 0.033 |
| | (0.0177) | (0.0266) | (0.0319) |
| Hh head working | 0.897*** | 1.147*** | 1.359*** |
| | (0.0250) | (0.0453) | (0.1091) |
| Hh head spouse working | 0.373*** | 0.490*** | -0.006 |
| | (0.0215) | (0.0335) | (0.0394) |
| Other income recipients | 0.267*** | 0.340*** | 0.279*** |
| | (0.0261) | (0.0384) | (0.0477) |
| Aboriginal | -0.113*** | -0.247*** | -0.242*** |
| | (0.0385) | (0.0595) | (0.0706) |
| Large city | 0.122*** | 0.172*** | 0.210*** |
| | (0.0207) | (0.0316) | (0.0377) |
| Victoria | -0.044* | -0.031 | -0.033 |
| | (0.0245) | (0.0369) | (0.0437) |
| Queensland | -0.056*** | -0.089*** | -0.084*** |
| | (0.0104) | (0.0156) | (0.0184) |
| South Aus | -0.105*** | -0.194*** | -0.151** |
| | (0.0356) | (0.0560) | (0.0672) |
| Western Aus | -0.042 | -0.047 | -0.029 |
| | (0.0350) | (0.0527) | (0.0623) |
| Tasmania | -0.005 | -0.131 | -0.085 |
| | (0.0584) | (0.0935) | (0.1095) |
| Northern Territory | 0.116 | 0.277 | 0.167 |
| | (0.1479) | (0.2247) | (0.2533) |
| Aus Capital Territory | 0.097 | 0.169 | 0.283** |
| | (0.0715) | (0.1032) | (0.1204) |
| Constant | 9.520*** | 9.394*** | 8.862*** |
| | (0.0314) | (0.0601) | (0.1177) |
| Observations | 16,667 | 15,339 | 13,930 |
| R-squared | 0.531 | 0.369 | 0.159 |

Notes: Table 2b presents the estimated OLS regression coefficients for each of the income measures as the dependent variable in each column. Coefficients on the interaction terms are omitted to save space. Standard errors are in parentheses. *** significant at the 1% level; ** significant at the 5% level; * significant at the 10% level.

The regression results for the determinants of consumption are, for the most part, as expected. The youngest cohort, households whose head was born in the 1970s, consumes significantly less across all consumption measures than the oldest cohort, those born in the 1940s. However, households whose head was born in the 1950s or 1960s in most cases consume more than the omitted category, although not statistically so. Households whose head completed a graduate degree consume more across all consumption measures, with the greatest difference in durables and the least in food. Larger households consume more, but after controlling for household size, having children in the family lowers consumption.

Families with three or more children consume less than families with one or two children, which could be explained by lower-income families more likely to rear more children. Households with the head and/or head's spouse working, as well as those with other income recipients, have significantly greater consumption. Surprisingly, aboriginals and Torres Strait Islanders in the HILDA sample are no less likely to consume than other ethnic groups. In fact, the coefficients on all consumption measures except durables are positive although not statistically significant. However, the sample size for this group is small (about 1 percent of the entire sample, see Table 1). Finally, households in large cities spend more on food and nondurables, but once durables are accounted for this difference is not statistically significant.

The results are somewhat different when income is the dependent variable. Compared to the oldest cohort in the sample, all other cohorts earn significantly less net income but significantly more male labor income. This is in part driven by lower female earnings for the older cohort, but evidence also suggests taxes are lower or transfers are greater for this group. Education and being resident in a large city is positively correlated with all income measures, as is head and/or head's spouse working as well as those households with other income recipients. Yet having a spouse work has no significant impact on male labor earnings. Although not the case for consumption, aboriginals and Torres Strait Islanders have significantly lower income, particularly labor income, showing that taxes and transfers are important.

One important caveat is that these regressions may remove changes that are unexpected by the individual. Blundell et al. (2008) acknowledge that this might change the relative degree of persistence in the remaining shocks. One example they give is that removing the effect of education-time on income and consumption (they interacted education with year dummies) could also remove the increase in inequality due to changing education premiums. The authors claim, however, that this should not affect the consumption-smoothing parameters.

In disagreement with the authors' claim, these regressions may in fact be removing the degree of consumption smoothing available to households. The shock is identified as unexplained income growth, and the explained growth in income takes into account labor supply decisions of households (for example, whether there are additional income recipients other than the household head). But labor supply (as well as savings) is an important margin of adjustment to absorb idiosyncratic shocks and contains substantial information about the degree of consumption smoothing available to individuals. And the main component of income is labor earnings. By treating these margins of adjustment in the regressions as expected changes, one is arguably removing the level of consumption smoothing that is being captured in the model.[11]

## b. The autocovariance of consumption and income

This section discusses underlying trends in (as well as the relationship between) unexplained income and consumption growth of Australian households by calculating moments within (and between) the series. This analysis is presented for the whole sample using the baseline income measure of net household income and baseline consumption measure of nondurables.

Tables 3-5 present several moments of the income and consumption processes.[12] The moments in Table 3 include the variance of the unexplained income growth, $\mathrm{var}(\Delta y_t)$, the first-order autocovariances, $\mathrm{cov}(\Delta y_{t+1}, \Delta y_t)$, and the second-order autocovariances, $\mathrm{cov}(\Delta y_{t+2}, \Delta y_t)$. Estimates are reported for each year between 2002 and 2009. Table 4 repeats the exercise for the consumption process between 2006 and 2009 (earlier moments are missing as nondurable consumption data was not collected prior to 2005). Table 5 presents estimates of contemporaneous and lagged consumption-income covariances between 2005 and 2009.

[11] One alternative would be to allow agents to adjust labor supply in response to idiosyncratic shocks using a general-equilibrium model. Alternative, it would be possible to instrument income shocks with wage shocks.
[12] All of the moments of unexplained income and consumption growth are found in the matrix $m$ and their variances in $V$.

The moments of the income process are informative about shifts in the income distribution as well as the transitory or permanent nature of such shifts. Although dropping between 2002 and 2003, the variance of unexplained income growth of Australian households remained relatively steady between 2003 and 2008 before increasing in 2009 by over 10 percent, suggesting that a shift in the income distribution of Australian households occurred between 2008 and 2009. Similar results are found in Figure 2 above that plots the variance of log net household income. The absolute values of the first-order autocovariances also show little deviation during the sample period, albeit with a small decline in 2003. Second-order autocovariances (and higher) are informative about the presence of serial correlation in the transitory income component. These moments are small, suggesting an MA(0) process in the transitory income component may be more appropriate than an MA(1) process (where an MA( 0 ) process in the transitory income component along with a serially uncorrelated permanent component would suggest an MA(1) process for income growth). However, to be consistent with Blundell et al. (2008), the estimation strategy still allows for MA(1) serial correlation in the transitory component.

**Table 3: The Autocovariance Matrix of Income Growth**

| Year | $\mathrm{var}(\Delta y_t)$ | $\mathrm{cov}(\Delta y_{t+1}, \Delta y_t)$ | $\mathrm{cov}(\Delta y_{t+2}, \Delta y_t)$ |
|---|---|---|---|
| 2002 | 0.0979 | -0.0366 | -0.0048 |
| | (0.0255) | (0.0171) | (0.0122) |
| 2003 | 0.0878 | -0.0243 | -0.0030 |
| | (0.0223) | (0.0139) | (0.0127) |
| 2004 | 0.0850 | -0.0311 | -0.0073 |
| | (0.0256) | (0.0162) | (0.0131) |
| 2005 | 0.0895 | -0.0303 | -0.0003 |
| | (0.0248) | (0.0167) | (0.0122) |
| 2006 | 0.0927 | -0.0275 | -0.0015 |
| | (0.0239) | (0.0147) | (0.0118) |
| 2007 | 0.0918 | -0.0317 | -0.0069 |
| | (0.0253) | (0.0148) | (0.0145) |
| 2008 | 0.0903 | -0.0309 | n.a. |
| | (0.0234) | (0.0186) | |
| 2009 | 0.1055 | n.a. | n.a. |
| | (0.0270) | | |

Notes: Table 3 presents, for each year, the variance of unexplained net household income growth (income growth minus its predictable components) in the first column, the first-order autocorvariance in the second column, and the second-order autocovariance in the third column. Standard errors are in parentheses.

Table 4 shows that unexplained consumption growth increased during the sample period, given by the magnitude of the variance increasing by about 20 percent between 2006 and 2009. This is consistent with Figure 1 above that plots the variance of log net household consumption. However, this difference between the estimates for 2006 and 2009 (equal to 0.0126) is not statistically significant, given by a p-value of 0.7060. If consumption indeed follows a martingale process, the variance of consumption growth should capture the impact of shocks to income. At first glance, this suggests there may be low consumption smoothing against income shocks occurring during the sample period. However, under this assumption, first-order autocovariances should also be zero. Instead, these are quite high, and one explanation could be that measurement error in the sample may be a problem. The absolute value of the first-order autocovriance of consumption growth should be a good estimate of the variance of the imputation error. To a first approximation, the variance of consumption growth that is not contaminated by error can be obtained by subtracting twice the (absolute value of) first-order autocovariances from the variance of consumption. The result would be minimal variance in unexplained consumption growth, suggesting instead that income shocks are not passing through to consumption. Second-order autocovariances are small.

**Table 4: The Autocovariance Matrix of Consumption Growth**

| Year | $\mathbf{var}(\Delta c_t)$ | $\mathbf{cov}(\Delta c_{t+1}, \Delta c_t)$ | $\mathbf{cov}(\Delta c_{t+2}, \Delta c_t)$ |
|---|---|---|---|
| 2006 | 0.1088 | -0.0596 | -0.0059 |
| | (0.0244) | (0.0192) | (0.0115) |
| 2007 | 0.1250 | -0.0446 | -0.0025 |
| | (0.0281) | (0.0154) | (0.0125) |
| 2008 | 0.1131 | -0.0527 | n.a. |
| | (0.0222) | (0.0148) | |
| 2009 | 0.1214 | n.a. | n.a. |
| | (0.0228) | | |

Notes: Table 4 presents, for each year, the variance of unexplained nondurable consumption growth (consumption growth minus its predictable components) in the first column, the first-order autocorvariance in the second column, and the second-order autocovariance in the third column. Standard errors are in parentheses.

Table 5 further supports this evidence by looking at moments in unexplained consumption and income growth in Australian households between 2005 and 2009. In fact, the contemporaneous covariances between income and consumption growth are small and declined between 2007 and 2009. While in 2006 and 2007 there is a strongly significant relationship between income and consumption, the relationship between income and consumption is not statistically different from zero at the conventional significance level for 2008 and 2009, at least at the household level. This is interesting to note given that, at least in 2009, Australia was significantly affected by the global financial crisis. The covariance between current income growth and future consumption growth should reflect the extent of consumption smoothing with respect to transitory shocks. Given these numbers are near to zero, this suggests there may be full consumption smoothing of transitory shocks. In addition, if shocks to income were known to the consumer beforehand, consumption should adjust prior to the shock occurring. This should show up in a positive and economically-significant covariance between changes in consumption and future income. Instead, the estimates of the covariance between current income growth and future consumption growth are also close to zero.

**Table 5: The Consumption-Income Growth Covariance Matrix**

| Year | $\mathrm{cov}(\Delta y_t, \Delta c_t)$ | $\mathrm{cov}(\Delta y_{t+1}, \Delta c_t)$ | $\mathrm{cov}(\Delta y_t, \Delta c_{t+1})$ |
|---|---|---|---|
| 2005 | n.a. | n.a. | 0.0004 |
| | | | (0.0100) |
| 2006 | 0.0038 | -0.0017 | -0.0012 |
| | (0.0011) | (0.0011) | (0.0012) |
| 2007 | 0.0047 | -0.0012 | 0.0004 |
| | (0.0013) | (0.0013) | (0.0012) |
| 2008 | 0.0011 | 0.0015 | 0.0007 |
| | (0.0012) | (0.0014) | (0.0010) |
| 2009 | 0.0016 | n.a. | n.a. |
| | (0.0012) | | |

Notes: Table 5 presents, for each year, the contemporaneous covariance of unexplained net household income growth (income growth minus its predictable components) and unexplained nondurable consumption growth (consumption growth minus its predictable components) in the first column, the covariance of future income growth and current consumption growth in the second column, and the covariance of current income growth and future consumption growth in the third column. Standard errors are in parentheses.

The moments presented in Tables 3-5 are, in general, precisely measured and statistically significant. Exceptions are the second-order autocovariances, which are also close to zero in magnitude. Another exception is the insignificant autocovariance between current consumption growth and future income growth. If households had advanced information of income shocks, future earnings growth should be correlated with current consumption growth, but this correlation in the data is not significant. This is important to note because advanced information would invalidate the identification strategy and bias the consumption-smoothing parameters.

This comment is related to the critique of the structural estimation approach that imposes restrictions on the stochastic income process faced by consumers. The model's identification strategy cannot statistically separate what is a shock from what is an anticipated event by the household, and instead assumes that the innovation process for income represents the arrival of new information to the household. That is, any shock is not foreseeable.

In reality, the individual may have advanced information about the evolution of future income (which is not observed by the econometrician). If this were the case, then this information should have already been incorporated into consumption choices. Consumption would change immediately upon this information being made available, thus reducing the magnitude of the change in the period of the innovation (given that consumers have preferences over smooth consumption profiles). As such, consumption would appear to react little to innovations in income, simply because they are anticipated by the agent and have already been incorporated into consumption choices. Thus the true consumption-smoothing parameter would be biased downward, that is, the econometrician would call consumption smoothing what is in part advanced information.

In general, it is hard to separate advanced information from consumption smoothing, which requires further information on whether the changes were or were not anticipated. Kaufmann and Pistaferri (2009) offer one solution to this identification problem by allowing a degree of advanced information in the income generating process. An alternative strategy for identifying the predictable component of earnings would be to explore survey questions, available in some datasets, where households are asked to report a probability distribution over changes in earnings in the next calendar year (Heathcote et al. 2009). More detailed data on private transfers and individual portfolios might also help in discriminating between insurability and forecastability.

## c. Baseline results

The baseline results of the DWMD estimations focus on the variances of the permanent and transitory income shocks, from here on denoted $\sigma_{\zeta}^2$ and $\sigma_{\epsilon}^2$ , respectively, and the consumption-smoothing parameters. These variances are allowed to vary across years but the consumption-smoothing parameters are assumed to be stationary over the sample period. An MA(1) process for the transitory component of income is assumed, $v_{i,t} = \epsilon_{i,t} + \theta\epsilon_{i,t-1}$, and the estimate of the MA(1) parameter $\theta$ is presented. In addition, the estimate of the variance of innovations to consumption growth independent of income growth, from here on denoted $\sigma_{\xi}^2$, is presented. Although included in the model, the results of the variances of consumption measurement error (which are also allowed to vary across years) are excluded. Finally, results are obtained for the total sample as well as among different subgroups of the population, including by education (graduate degree versus no degree) and by cohort (born in the 1950s versus born in the 1960s). Table 6 presents the results.

**Table 6: Minimum Distance Parameter Estimates**

| | | Whole | No college | College | Born 1950s | Born 1960s |
|---|---|---|---|---|---|---|
| $\sigma_{\zeta}^2$ (variance of perm. shock) | 2002-04 | 0.0183 | 0.0230 | 0.0157 | 0.0242 | 0.0168 |
| | | (0.0014) | (0.0015) | (0.0010) | (0.0012) | (0.0012) |
| | 2005 | 0.0195 | 0.0217 | 0.0089 | 0.0262 | 0.0169 |
| | | (0.0019) | (0.0020) | (0.0014) | (0.0017) | (0.0018) |
| | 2006 | 0.0283 | 0.0339 | 0.0209 | 0.0475 | 0.0140 |
| | | (0.0020) | (0.0020) | (0.0020) | (0.0020) | (0.0018) |
| | 2007-09 | 0.0263 | 0.0331 | 0.0148 | 0.0338 | 0.0195 |
| | | (0.0014) | (0.0014) | (0.0011) | (0.0013) | (0.0013) |
| $\sigma_{\epsilon}^2$ (variance of trans. shock) | 2002 | 0.0446 | 0.0449 | 0.0383 | 0.0308 | 0.0393 |
| | | (0.0019) | (0.0019) | (0.0013) | (0.0012) | (0.0016) |
| | 2003 | 0.0328 | 0.0277 | 0.0353 | 0.0341 | 0.0339 |
| | | (0.0016) | (0.0016) | (0.0013) | (0.0016) | (0.0017) |
| | 2004 | 0.0394 | 0.0409 | 0.0371 | 0.0270 | 0.0461 |
| | | (0.0019) | (0.0019) | (0.0018) | (0.0014) | (0.0023) |
| | 2005 | 0.0379 | 0.0352 | 0.0404 | 0.0309 | 0.0387 |
| | | (0.0017) | (0.0018) | (0.0012) | (0.0016) | (0.0018) |
| | 2006 | 0.0338 | 0.0274 | 0.0443 | 0.0103 | 0.0334 |
| | | (0.0017) | (0.0016) | (0.0018) | (0.0010) | (0.0018) |
| | 2007-09 | 0.0394 | 0.0333 | 0.0469 | 0.0330 | 0.0334 |
| | | (0.0015) | (0.0014) | (0.0014) | (0.0011) | (0.0013) |
| $\theta$ (serial correlation of trans. shock) | | 0.1069 | 0.0486 | 0.1708 | -0.0214 | 0.1288 |
| | | (0.0127) | (0.0146) | (0.0077) | (0.0169) | (0.0144) |
| $\sigma_{\xi}^2$ (variance of consumption innovation) | | 0.0111 | 0.0085 | 0.0187 | 0.0070 | 0.0055 |
| | | (0.0012) | (0.0012) | (0.0013) | (0.0008) | (0.0012) |
| $\phi$ (consumption smoothing perm. shock) | | 0.0917 | 0.0960 | 0.0882 | 0.0831 | 0.1850 |
| | | (0.0284) | (0.0244) | (0.0443) | (0.0164) | (0.0351) |
| $\psi$ (consumption smoothing trans. shock) | | 0.0102 | -0.0004 | 0.0193 | -0.0108 | 0.0003 |
| | | (0.0231) | (0.0255) | (0.0202) | (0.0228) | (0.0212) |

Notes: Table 6 presents the diagonally weighted minimum distance parameter estimates for five samples: the whole sample in the first column, households whose head does not have a graduate degree in the second column, households whose head has a graduate degree in the third column, households whose head was born in the 1950s in the fourth column, and households whose head was born in the 1960s in the fifth column. Standard errors are in parentheses.

For the entire sample of Australian households, nearly full consumption smoothing exists against transitory shocks, as indicated by the consumption-smoothing parameter estimate of $\psi$

of 0.01. In addition, this estimate is not statistically different from zero. Less consumption smoothing exists against permanent shocks, with a parameter estimate of $\phi$ of just above 0.09, which is statistically different from zero. These empirical results suggest that a 10 percent permanent income shock induces about a 1 percent change in consumption. This further suggests that for the average Australian household during the sample period, the marginal propensity to consume out of a permanent AUD 1 increase in income is about 30 cents. It is interesting to note that this estimate is substantially higher than what previous literature had found for Australia during the 1980s and 1990s (between 3 and 10 cents). However, Australian households still achieve a high degree of consumption smoothing even against highly persistent shocks, particularly when compared to households in the United States. Blundell et al. (2008) estimate $\psi$ to be 0.05, not much higher than this study's estimate for Australia, and also conclude that households have full consumption smoothing against transitory income shocks. However, the estimate of $\phi$ is significantly higher than in Australia at 0.64.

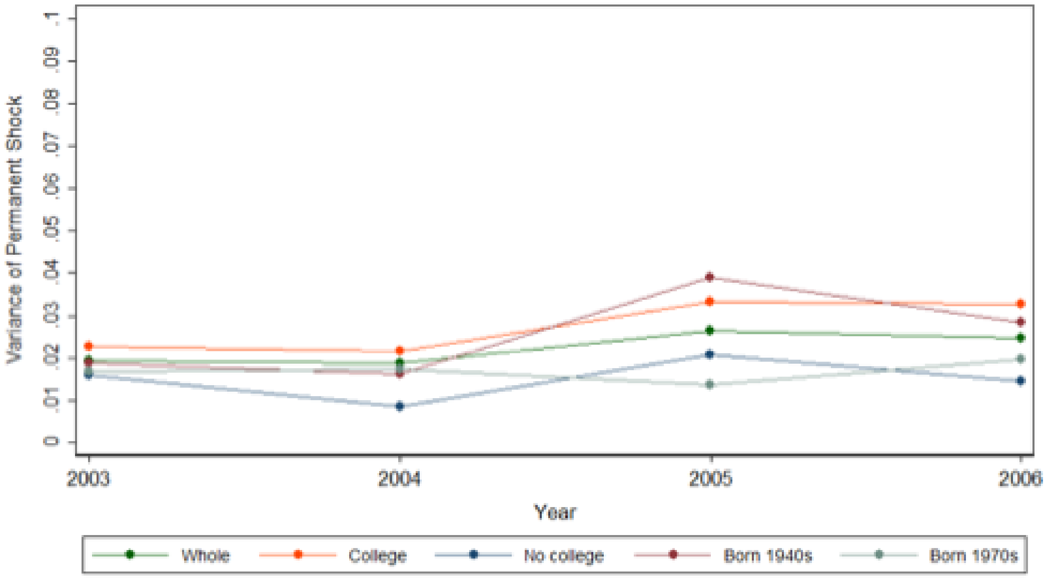
Overall the results change little for different subgroups of the sample. Households without a graduate diploma have slightly less consumption smoothing against permanent shocks than those with a graduate diploma. The younger cohort, with household heads born in the 1960s, show lower ability to smooth consumption against permanent shocks than the older cohort, with household heads born in the 1950s. However, for all these subgroups, full consumption smoothing exists against transitory shocks. Again, the consumption-smoothing parameter of permanent shocks is precisely estimated whereas of transitory shocks is not statistically different from zero. Similar trends are found in the United States, although the magnitudes of the estimates are different. For example, permanent income shocks pass almost fully through to consumption for the sample of American households whose head has no college education, substantially higher than those with a college education (estimated at 0.94 versus 0.42, respectively). In addition, the MA parameter for the transitory shock is positive but small.

While the evidence on the consumption-smoothing parameter to transitory shocks is consistent with the standard permanent income hypothesis (households have full consumption smoothing to transitory shocks), the evidence on the consumption-smoothing parameter to permanent shocks is instead inconsistent. The model predicts that permanent changes in income should have significant effects on consumption. Rather, the baseline results for Australia accord well with the complete markets hypothesis, which predicts full consumption smoothing to idiosyncratic income shocks (both transitory and permanent).

In addition, the result that permanent shocks are smoothed to a greater extent by older cohorts than younger cohorts offers some support for consumption smoothing through precautionary savings. The model predicts that all individuals will use precautionary savings to completely smooth consumption against transitory shocks. But for younger cohorts, whose value of current financial assets is small relative to remaining future labor income, permanent shocks pass through completely to consumption. Although there is evidence in Australia that permanent shocks for the older cohort are smoothed to a greater extent than for younger cohorts, younger cohorts are still able to partially smooth consumption against permanent shocks.

The variance of the transitory shock changed little between 2002 and 2009, while the variance of the permanent shock increased nearly 50 percent between the beginning and end of the sample period (plotted over time in Figure 3 for the whole sample as well as each of the subgroups). The results also show that households have less consumption smoothing against permanent versus transitory shocks. Together these results provide one explanation for the observed increase in nondurable consumption inequality of Australian households in the later part of the sample period: idiosyncratic income shocks were becoming more persistent, which households are less able to smooth consumption against. However, other explanations are possible that are outside the scope of this study to consider. As pointed out in Equation (3), a lower degree of consumption smoothing (implied by higher values for $\phi$ or $\psi$) could also explain such a finding. Unfortunately there are not enough time periods in the sample to allow these parameters to change with time.

**Figure 3: Variance of Permanent Shocks**




Notes: Figure 3 plots the evolution of the diagonally weighted minimum distance parameter estimates for the variance of the permanent shock from 2003 to 2006 for five samples: the whole sample, households whose head does not have a graduate degree, households whose head has a graduate degree, households whose head was born in the 1940s, and households whose head was born in the 1970s.

## d. Results by consumption and income measures and population groups

This section extends the analysis to explore how these results change for alternative definitions of consumption and income as well as for different subgroups of the population. Table 7 presents the consumption-smoothing parameter estimates for the whole sample using each of the different measures of consumption. The results for food as well as the nondurable measure containing durable services mimic closely those obtained using the baseline nondurable consumption measure.

However, when the definition contains durables, the estimate of the consumption-smoothing parameter of permanent shocks nearly doubles. That is, expenditures on durable goods react much more to unexplained income changes than expenditures on nondurable goods. This suggests that durables may act as an intertemporal consumption-smoothing mechanism. The timing of purchases or replacement of non-fully collateralized durables could be used to smooth nondurable consumption in the face of income shocks: households may forgo nondurable consumption goods today to purchase durable goods that in the future can be converted back to nondurables. This implies that less evidence for consumption smoothing should be found with a measure of consumption that includes durables, which appears to be the case for permanent shocks to income but not transitory shocks.[13] In contrast, in the United States the consumption-smoothing parameters for both transitory shocks and permanent shocks increased for low-wealth households.

**Table 7: Consumption Measures**

| **Consumption:**<br>**Income:**<br>**Sample:** | **Nondurable**<br>**Net hh income**<br>**Whole** | **Food**<br>**Net hh income**<br>**Whole** | **Durable services**<br>**Net hh income**<br>**Whole** | **Durables**<br>**Net hh income**<br>**Whole** |
|---|---|---|---|---|
| $\phi$ | 0.0917 | 0.0929 | 0.0944 | 0.1639 |
| (consumption smoothing perm. shock) | (0.0284) | (0.0183) | (0.0319) | (0.0449) |
| $\psi$ | 0.0102 | -0.0051 | -0.0044 | 0.0008 |
| (consumption smoothing trans. shock) | (0.0231) | (0.0158) | (0.0265) | (0.0339) |

Notes: Table 7 presents the diagonally weighted minimum distance parameter estimates of consumption smoothing to permanent shocks and transitory shocks for the whole sample using net household income and each of the different measures of consumption (nondurable in the first column, food in the second column, durable services in the third column, durable in the fourth column). Standard errors are in parentheses.

Blundell et al. (2008) found taxes, transfers, and family labor supply to play important roles in insuring permanent shocks. The extent to which these act as consumption-smoothing mechanisms in Australia is explored in Table 8, which presents the consumption-smoothing parameter estimates for the whole sample using each of the different measures of income. The reduction in the consumption-smoothing parameter for permanent shocks indicates that taxes and transfers as well as female labor supply provide consumption smoothing for households against permanent income shocks. While the nondurable consumption measure still incorporates the consumption smoothing value of taxes and transfers and female labor supply, these new measures of income no longer do.

[13] Cerletti and Pijoan-Mas (2012) find that borrowing constraints lead to a substitution between durable and non-durable goods upon arrival of an unexpected income change, and this substitution can bias the measure of the extent of consumption smoothing when using non-durables. Therefore, true consumption smoothing against permanent shocks is larger than what is typically measured by the response of non-durable expenditure only.

**Table 8: Income Measures**

| Consumption:<br>Income:<br>Sample: | Nondurable<br>Net hh income<br>Whole | Nondurable<br>Hh labor<br>Whole | Nondurable<br>Male labor<br>Whole |
|---|---|---|---|
| $\phi$ | 0.0917 | 0.0652 | 0.0300 |
| (consumption smoothing perm. shock) | (0.0284) | (0.0025) | (0.0100) |
| $\psi$ | 0.0102 | 0.0049 | -0.0491 |
| (consumption smoothing trans. shock) | (0.0231) | (0.0105) | (0.0460) |

Notes: Table 8 presents the diagonally weighted minimum distance parameter estimates of consumption smoothing to permanent shocks and transitory shocks for the whole sample using nondurable consumption and each of the different measures of income (net household in the first column, household labor in the second column, male labor in the third column). Standard errors are in parentheses.

Finally, Table 9 explores whether the results are heterogenous across different population groups, including low-income households and households whose head is aboriginal or Torres Straight Islander. Following Blundell et al. (2008), the wealth of each household $i$ is calculated the first year $t$ that the household is observed in the sample as: $\text{asset income}_{i,t}/r_t + \text{housing income}_{i,t}$, where $r_t$ is the interest rate on a 2-year government bond averaged across 12 months of the year reported by the Reserve Bank of Australia.[14] A low-wealth household is a household in the bottom 20 percent of the wealth distribution in the first year $t$ that the household is observed in the sample. While in the United States Blundell et al. (2008) find a significant impact of transitory shocks on consumption for low-wealth households, the findings here suggest that households in Australia are able to fully smooth consumption against transitory shocks. What is different, however, is that low-income households are much less able to smooth consumption against permanent shocks. The finding that wealthier households smooth consumption better than low-income households is also consistent with evidence that time preference is higher for these individuals who prefer immediate gratification. Finally, households whose head is aboriginal or Torres Straight Islander appear to be more able to smooth consumption to both permanent and transitory income shocks.

[14] As Blundell et al. (2008) point out, given that the level of wealth in the initial period is pre-determined with respect to consumption growth decisions thereafter, the corresponding sample stratification does not suffer from endogeneity problems.

**Table 9: Population Groups**

| Consumption:<br>Income:<br>Sample: | Nondurable<br>Net hh income<br>Whole | Nondurable<br>Net hh income<br>Low wealth | Nondurable<br>Net hh income<br>Non-low wealth | Durable<br>Net hh income<br>Low wealth | Nondurable<br>Net hh income<br>Aboriginal |
|---|---|---|---|---|---|
| $\phi$ | 0.0917 | 0.2958 | 0.0878 | 0.5024 | 0.0149 |
| (consumption smoothing perm. shock) | (0.0284) | (0.0186) | (0.0265) | (0.0434) | (0.0030) |
| $\psi$ | 0.0102 | 0.0334 | -0.0025 | 0.0011 | 0.0060 |
| (consumption smoothing trans. shock) | (0.0231) | (0.0398) | (0.0231) | (0.0542) | (0.0198) |

Notes: Table 9 presents the diagonally weighted minimum distance parameter estimates of consumption smoothing to permanent shocks and transitory shocks for different samples (whole in the first column, low wealth in the second and fourth columns, non-low wealth in the third column, aboriginal or Torres Straight Islander in the fifth column) using net household income and different measures of consumption (nondurable in all columns except durable in the fourth column). Standard errors are in parentheses.

# 6. Concluding Remarks

This study examined the link between individual-specific changes in income and changes in consumption of Australian households over the period 2001-2009 using the HILDA dataset. In particular, it estimated the degree of transmission of permanent as well as transitory idiosyncratic income shocks to consumption following the methodology of Blundell et al. (2008). The degree of these transmissions, called the "partial consumption-smoothing parameters" of permanent and transitory shocks, were identified from the authors' permanent-transitory model that characterizes the processes of unexplained income growth and consumption growth. The model's parameters were then estimated using Australian household-level panel data. This analysis was disaggregated by education, cohort of birth, and wealth to examine whether heterogeneity exists in the degree of consumption smoothing across different population subgroups. In addition, it empirically analyzed the mechanisms behind the degree of consumption smoothing found in the data, in particular the role of durable purchases, female labor supply, and taxes and transfers.

For the entire sample of Australian households, nearly full consumption smoothing exists against transitory shocks that are specific to the individual. Although less consumption smoothing exists against permanent shocks, Australian households still achieve a high degree of consumption smoothing against highly persistent shocks, particularly when compared to households in the United States. The study's empirical results suggest that a 10 percent permanent income shock induces about a 1 percent change in consumption. Although Blundell et al.'s (2008) estimate of the consumption-smoothing parameter of transitory shocks is similar in magnitude to that found in Australia, the estimate of the consumption-smoothing parameter of permanent shocks is significantly higher. In the United States, a 10 percent permanent income shock was found to induce about a 6 percent change in consumption. This suggests that for the average Australian household during the sample period, the marginal propensity to consume out of a permanent dollar increase in income is about 30 cents. In addition, there is reason to believe Blundell et al.'s (2008) estimates for the United States are biased downwards due to measurement error in imputed consumption. The baseline results for Australia accord well with the complete markets hypothesis, which predicts full consumption smoothing to idiosyncratic income shocks.

These results of the marginal propensity to consume are above those found in previous Australian studies. Using a panel of Australian states for 1984-2001, Dvornak and Kohler (2007) find that changes in housing and stock market wealth have a significant effect on consumption expenditure: a permanent AUD 1 increase in stock market wealth increases annual consumption by about 9 cents, and the same increase in housing wealth increases annual consumption by about 3 cents. Tan and Voss (2003) have estimated that during the period 1988-1999, annual consumption increased by 4 cents in response to an AUD 1 increase in wealth. These estimates are lower than those found in this study. Using household-level data, Berger-Thomson et al. (2010) estimate the marginal propensity to consume in Australia due to two policy changes: income tax rates and lump-sum transfers (the baby bonus). While the marginal propensity to consume is found to be unity for tax cuts, it is 0.1 for lump-sum transfers.

When the analysis was disaggregated across different population subgroups, there was some support for consumption smoothing through precautionary savings. Permanent shocks are smoothed to a greater extent by older cohorts than younger cohorts (although younger cohorts are still able to partially smooth consumption against permanent shocks). Similar to the United States, low-wealth households in Australia are less able to smooth consumption against permanent shocks. Durable purchases, female labor supply, and taxes and transfers were all found to act as consumption-smoothing mechanisms.

The joint evolution of consumption and income inequality of Australian households during this time period was also examined to determine whether changes in the persistence of income shocks have affected the evolution of income and consumption inequality. Consumption inequality (measured as the variance of log household nondurable consumption) of Australian households increased slightly in the later part of the sample period. The results showed that the variance of the transitory shock changed little between 2002 and 2009, while the variance of the permanent shock increased nearly 50 percent between the beginning and end of the sample period. The results also showed that households have less consumption smoothing against permanent versus transitory shocks. Together these results provide one explanation for the observed increase in consumption inequality: idiosyncratic income shocks were becoming more persistent, which households are less able to smooth consumption against.

# Works Cited


Abowd, J. and D. Card (1989). "On the Covariance Structure of Earnings and Hours Changes," *Econometrica*, 57(2): 411-45.

Berger-Thomson, L., E. Chung, and R. McKibbin (2010). "Estimating Marginal Propensities to Consume in Australia Using Micro Data," *The Economic Record*, 86: 49-60.

Blundell, R., L. Pistaferri, and I. Preston (2008). "Consumption Inequality and Partial Insurance," *American Economic Review*, 98(5): 1887-1921.

Blundell, R., and I. Preston (1998). "Consumption Inequality and Income Uncertainty," *Quarterly Journal of Economics*, 113(2): 603-40.

Casado, J. M. (2011). "From Income to Consumption: Measuring households partial insurance," *Empirical Economics*, 40(2): 471-95.

Cerletti, E. A., and J. Pijoan-Mas (2012). "Durable Goods, Borrowing Constraints and Insurance," CEMFI Working Paper No. 1206, Center for Monetary and Financial Studies, Madrid.

Dvornak, N., and M. Kohler (2007). "Housing Wealth, Stock Market Wealth and Consumption: A panel analysis for Australia," *The Economic Record*, 83(261): 117-30.

Guvenen, F., and A. A. Smith (2008). "Inferring Labor Income Risk from Economic Choices: An Indirect Inference Approach," unpublished.

Hall, R. E., and F. S. Mishkin (1982). "The Sensitivity of Consumption to Transitory Income: Estimates from Panel Data on Households," *Econometrica*, 50(2): 461-81.

Hayashi, F., J. Altonji, and L. Kotlikoff (1996). “Risk-Sharing Between and Within Families,” *Econometrica*, 64(2): 261-94.

Heathcote, J., K. Storesletten, and G. L. Violante (2009). “Consumption and Labor Supply with Partial Insurance: An Analytical Framework,” Working Paper No. 15257, National Bureau of Economic Research, Cambridge, MA.

Holloway, J. C. (1992). “The Rational Expectations Permanent Income Hypothesis: Does aggregate consumption in Australia follow a random walk?,” unpublished.

Johnson, P. (1983). “Life-Cycle Consumption under Rational Expectations: Some Australian evidence,” *The Economic Record*, 59(4): 345-50.

Kaplan, G., and G. Violante (2010). “How Much Consumption Insurance Beyond Self-Insurance?,” *American Economic Journal: Macroeconomics*, 2(4): 53-87.

Kaufmann, K., and L. Pistaferri (2009). “Disentangling Consumption Smoothing and Information in Intertemporal Consumption Choices,” *American Economic Review*, 99(2): 387-92.

McCurdy, T. H. (1982). “Dynamics Aspect of Earnings Mobility,” *Journal of Econometrics*, 18: 83-114.

MacDonald, R., and C. Kearney (1990). “Consumption, Cointegration and Rational Expectations: Some Australian evidence,” *Australian Economic Records,* 29(54): 40-52.

McKibbin, W., and A. Richards (1988). “Consumption and Permanent Income: The Australian case,” Research Discussion Paper No. 8808, Research Department, Reserve Bank of Australia, Sydney, Australia.

Meyer, B. D., and J. X. Sullivan (2013). "Consumption and Income Inequality and the Great Recession," *American Economic Review*, 103(3): 178-83.

Souleles, N. S., J. A. Parker, and D. S. Johnson (2006). "Household Expenditure and the Income Tax Rebates of 2001," *American Economic Review*, 96(5): 1589-610.

Primiceri, G. E., and T. van Rens (2009). "Heterogenous life-cycle profiles, income risk and consumption inequality," *Journal of Monetary Economics*, 56(1): 20-39.

Tan, A., and G. Voss (2003). "Consumption and Wealth in Australia," *The Economic Record*, 79(244): 39-56.

# Chapter 2

# Structural Reforms and Labor-Market Outcomes: International panel-data evidence

**Claire H. Hollweg**
**with Daniel Lederman**
**and Devashish Mitra**

# 1. Introduction and Motivation

The debate over the effects of structural reforms has been revived in recent years, in part because of the ongoing debt crises in European economies. As countries engage in fiscal stabilization through austerity policies, pressure has been mounting for debt-ridden economies to undertake structural reforms in conjunction with macroeconomic stabilization. The main argument in favor of structural reforms is that they tend to improve productivity and thus may be one of the few potential sources of growth in the context of fiscal austerity. From the viewpoint of developing countries, we have seen this movement before, when numerous economies undertook both fiscal stabilization and structural reforms throughout the 1980s and 1990s.

This study contributes to the relevant literature by focusing on the impact of structural reforms on macro-level labor-market outcomes. There is need for serious research on the impact of structural reforms on a comprehensive list of labor-market outcomes, which is what we try to do in this study. The outcomes that we study are the unemployment rate, the employment level, average wage index, labor force participation rates (overall and female) and the female employment level. We utilize publicly available data on labor-market outcomes as well as on the dates of reforms to assess how labor market trends were influenced by the advent of structural reforms within countries. We construct a structural reform index that proxies for the date when countries reached a threshold of broad reforms including macroeconomic stabilization, privatization, trade opening, as well as the end of interventionist states such as communism.

We first document the average trends (across countries) in our labor-market outcomes up to ten years on either side of each country's reform year. This is done by controlling for country fixed effects and in another specification additionally controlling for real GDP, the labor force participation rate and the working age population. On average we find that unemployment rates are higher after the reform than before the reform. In addition, we find that employment is higher on average and its trend is more positive after as compared to the early years before the reforms. Somewhat similar trends are also found with wages and labor force participation rates (overall and female). In the case of labor force participation rates, it is the trend that takes an upward turn after the reforms.

We next run fixed-effects ordinary least squares as well as instrumental variables regressions of our labor-market outcomes on a reform dummy variable, a time trend, square of the time trend and the same set of controls mentioned above. Focusing on wages, the Stolper-Samuelson theorem implies that in developing countries, which generally are abundant in unskilled labor, liberalization will increase the wage rate of unskilled workers. However, what are questioned then are the assumptions under which this theorem holds. One basic assumption that is questioned is free mobility of labor across sectors. When labor does not move across sectors (is sector-specific) and earns a higher wage in the protected, sophisticated manufacturing sectors, the fear is that removing protection will lower these workers' wage. It is also feared that the presence of adjustment costs could lead to higher unemployment after reforms at least in the short run, if not in the long run. While trade reforms are a very important component of structural reforms, other aspects of structural reforms include macroeconomic stabilization, privatization, and deregulation. Such reforms provide a better macroeconomic environment along with better economic incentives for everyone, leading to higher productivity and therefore higher wages and employment. However, moving to a better environment also involves short-run adjustments, during which we might see some adverse consequences on labor-market outcomes. The interpretation of our instrumental variable estimates on the effect of reforms on unemployment rates can range from inconclusive evidence to some support for the presence an unemployment-reducing effect of reforms.

The effects of our regressions for the employment level and the wage index are quite strong and conclusive in that, even in the presence of a strongly significant and positive time trend that we control for, reforms have a positive and significant impact on employment and wages. Because the labor force participation rate of the formal sector is included in the regressions and due to the weak results on unemployment after reforms, the results suggest that part of the increase in employment may be from informal workers. Perhaps the study has little to say about what would happen in Europe since most of the positive effects on both employment and labor force participation rates might be in part related to the formalization of previously informal workers. Although in the context of Greece, it is unclear that informality is not an issue.

The evidence on the effect of structural reforms on labor force participation rates (overall and female) is somewhat inconclusive even though we find that there is a positive time trend. Finally, there is some indication that structural reforms may have increased female employment. Because we have controlled for the time trend and real GDP, our results show the impact of structural reforms on labor-market outcomes beyond what happens through the impact on growth. Redistributive effects in favor of workers, along the lines of the Stolper-Samuelson effect, may be at work.

While the literature that finds evidence of a positive effect of structural reforms and greater openness on growth is well established, to our knowledge, there are only two major cross-country empirical studies that look at the impact of trade policy on unemployment rates.[15] One is the paper by Dutt, Mitra and Ranjan (2009) and the other is by Felbermayr, Prat and Schmerer (2011a).[16] Both papers show that countries that have less protectionist (more open)

[15] Included in the literature that finds evidence of structural reforms leading to growth is the seminal work of Sachs and Warner (1995). While the paper's measurement and estimation framework came under criticism, Wacziarg and Welch (2008) show that liberalization has, on average, robust positive effects on growth. See Baldwin (2003) for a survey of the literature.

[16] See also Hasan et al. (2012) for an intra-country study on cross-state and cross-industry variations in unemployment in the context of trade reforms.

trade policies have lower unemployment rates. This is true both without any controls and after controlling for other policies and institutions that have a more direct impact on labor markets. Dutt, Mitra and Ranjan (2009) also find that the short-run impact of trade reforms is an increase in the unemployment rate followed by a reduction in the long run to a lower steady-state unemployment rate. In addition, there are papers studying the role of other kinds of policies and institutions, namely labor market policies and institutions in the determination of unemployment rates. Scarpetta (1996) uses a panel of OECD countries to study the role of labor market policies and institutions on unemployment rates and the speed of adjustment of the labor market. Another paper looking at the role of labor market institutions on unemployment rates in the OECD is Nickell, Nunziata and Ochel (2005). Blanchard and Wolfers (2000) look at the interaction between macroeconomic shocks (such as shocks to total factor productivity and the interest rate) and labor market institutions (such as the degree of employment protection) in the determination of unemployment rates in Europe.

Our study differs from Dutt, Mitra and Ranjan (2009) and Felbermayr, Prat and Schmerer (2011a) along several dimensions. The first major and probably the main difference is that while these papers focus on the unemployment rate, our study looks at a multitude of labor-market outcomes. Secondly, unlike the previous studies, we include sample periods in our panel, such that for each country we cover a maximum of ten years before the reforms and a maximum of ten years after. While this approach can limit the number of observations, it gives us greater confidence in our results for the countries we are able to study. Thirdly, unlike the two papers mentioned above, our focus is not limited to trade policies. We look at the impact of structural reforms in general (of which trade reform is just one component) on labor-market outcomes. Finally, we are able to use credible instrumental variables for our structural reform variable. It is well known that structural reforms are endogenous to macroeconomic policies and conditions and therefore to unemployment rates and other labor-market outcomes. We draw on the political economy literature to come up with a fairly extensive list of instruments. The overidentification of the equations we estimate allows us to econometrically check the quality of instruments.

Although similar in spirit, there are reasons to expect our results of the impact of structural reforms on unemployment to differ from these papers' results of the impact of trade on unemployment. Dutt, Mitra and Ranjan (2009) and Felbermayr, Prat and Schmerer (2011a) empirically search for a causal relationship between openness and unemployment and both papers find that more open economies tend to have lower unemployment levels. In contrast, we find only some support for a causal impact of structural reforms on unemployment, at least within ten years of reform. Our broader focus on structural reforms rather than just trade policy along with the coverage of a longer time period (which spans symmetrically from ten years before to ten years after each reform) leads us to different results. The weakness however is that we do not go beyond ten years after the reform. Structural reforms involve short-run adjustments during which we might see some adverse consequences on labor-market outcomes. In fact, Dutt, Mitra and Ranjan (2009) do find the short-run impact of trade reforms to be an increase in the unemployment rate followed by a decrease in the long-run to a lower steady-state unemployment rate.

The remainder of the chapter proceeds as follows. Section 2 discusses theoretical predictions of the impacts of structural reforms on labor-market outcomes and why structural reforms are likely endogenous to these outcomes. Section 3 details the econometric methodology that accounts for this endogeneity. Section 4 presents the data and Section 5 the results. Section 6 concludes.

## 2. Theory

### a. The impact of structural reforms on labor-market outcomes

In this section, we will discuss possible theoretical predictions of the impact of structural reforms and liberalization on the unemployment rate, the level of employment, the wage rate and the labor force participation rate. We will not present just one model but will describe some prominent models in the literature and their predictions. Here we want to reiterate that our focus is not just limited to trade reforms. We look at the impact of structural reforms in general. However, it is important to note that trade reforms are a very important part of structural reforms and so a large part of our discussion will be about the impact of trade reforms on labor-market outcomes.

In particular, we highlight the gains from considering the impact of structural reforms and liberalization on additional labor market variables other than just the unemployment rate. For example, the unemployment rate does not capture displaced workers leaving the labor force rather than seeking alternative employment, which would be captured when considering employment and the labor force participation rate. Or considering the movement of wages alongside these variables may highlight the channels through which structural reforms impact these additional outcomes. Thus we expect to have a broader picture of the labor market implications of structural reforms by considering employment, labor force participation rates, and wages.

Krugman (1993) has argued that "…the level of employment is a macroeconomic issue, depending in the short run on aggregate demand and depending in the long run on the natural rate of unemployment, with microeconomic policies like tariffs having little net effect." Subsequent empirical work as well as theoretical developments have shown that this is not the case. For example, Trefler (2004) clearly provided conclusive evidence of a reduction in employment (and increase in the unemployment rate) in Canada arising from the short-run adjustment costs upon the signing and implementation of the NAFTA. Recent cross-country studies (Dutt, Mitra and Ranjan 2009, Felbermayr, Pratt and Schmerer 2011a) show how unemployment rates and trade protection are positively related.

In the context of this study, an important issue is that of micro- versus macro-level studies. While micro-level studies are very popular within and outside the academic economics profession, macro-level studies are not as well respected these days. However, we argue that broad questions of the sort we are dealing with in this study can only be answered through cross-country, macro-level panel regressions. Micro-level studies, which find that wages and/or employment go up in some sectors and fall in others and within a sector go up in some firms and fall in others in response to structural reforms, are not of great value to policy makers when they have to decide on broad policy reforms such as trade reforms, industrial deregulation, tax reforms, etc.

Although little empirical work exists on the aggregate labor-market effects of structural reforms, our study is related to other important research that estimates cross-country unemployment regressions. The literature is mainly concerned with the impact of macroeconomic shocks on labor market institutions, including, for example, Blanchard and Wolfers (2000) and Scarpetta (1996). (See Bassanini and Duval (2006, 2009) for a survey of this literature.) Other more recent examples include Nickell et al. (2005) and Boulhol (2008). Boulhol (2008) focuses on trade openness interacted with labor market institutions but does not address the endogeneity of this relationship. However, this literature has focused primarily on developed OECD countries. In addition, many of these studies are more about the long-run relationships than about short-run effects. On the theoretical side, however, there is a long list of papers showing the impact of trade policy on the structural unemployment rate.

Consider a two-sector Ricardian model with labor as the only factor of production. But, unlike a standard Ricardian model, suppose there are search frictions in the labor market. If we assume that the nature and extent of search frictions are the same in both sectors, trade liberalization leads to an increase in the value of the marginal product of labor (along with specialization in one of the two goods). Each worker will get a higher wage. For a given labor force participation rate, more vacancies are created relative to potential workers searching for jobs. This will make the labor market tighter and lower the unemployment rate in the long run, after short-run adjustments have taken place. How long it takes for these short-run adjustments to be completed is an empirical issue. Note that such a change will also lead to an increase in the labor force participation rate as the higher wage creates a greater incentive to look for a job. Thus, overall employment will increase.

From here, let us move to a two-factor, two-sector Heckscher-Ohlin framework. Let us call the two factors labor and capital. Let us assume that capital services are sold in a perfectly competitive market while there are search frictions in the labor market. A labor-abundant country has a comparative advantage in the labor-intensive good. Therefore, upon opening to trade, the structure of such an economy gets more specialized towards the labor-intensive good. This increases labor demand and the value of the marginal product of labor goes up, increasing the incentive for posting more vacancies. This, in turn, results in an increase in labor-market tightness and, therefore, an increase in the wage and the labor force participation rate and a reduction in the unemployment rate. It is easy to see that the results in the case of a capital-abundant economy will just be the opposite. Another possible productivity-driven, unemployment-reducing effect of trade is fully worked out in Felbermayr, Prat and Schmerer (2011). The effect of trade in that paper works through an interfirm labor reallocation effect in a one-sector model of search frictions with monopolistic competition, increasing returns to scale and heterogeneous firm productivity. After trade liberalization, the least productive firms are not able to survive the greater competition from more productive domestic and foreign firms. Also, the more productive firms grow at the expense of the other firms. The average productivity level therefore expands and, as a result, so does employment and wages. Unemployment falls upon trade liberalization. Helpman and Istkhoki (2009) construct a similar model but they have a second sector, which is one with homogeneous productivity, constant returns to scale and perfect competition. The impact of trade reforms on unemployment is ambiguous in that model. Examples of other models in which trade reforms have ambiguous effects on unemployment are Davidson, Martin and Matusz (1999) and Moore and Ranjan (2005). In Davidson, Martin and Matusz, the results depend on international and intersectoral differences in search frictions. On the other hand, in Moore and Ranjan (2005) the results are driven by the assumption that there are two types of labor (skilled and unskilled).

As mentioned earlier, other components of structural reforms that include macroeconomic stabilization, privatization, deregulation, etc., create a better economic environment and better economic incentives for businesses and workers. These improvements raise productivity and lead to better labor-market outcomes.

We next discuss the possible short-run effects of structural reforms on unemployment. In the short-run, labor would hardly be able to move from one sector to another. So we assume no intersectoral mobility of labor in the short run. If there is another factor such as capital, then even that factor would not be able to move so quickly. In addition, we could think of job destruction as endogenous, along the lines of Chapter 2 in Pissarides (2000). A firm-job pair starts at full productivity at the point of creation, but this is followed by a series of productivity shocks over time that are received by each firm at a Poisson arrival rate. The threshold productivity level for firm survival is the one at which the firm just breaks even, below which the firm-worker pair is destroyed. Since revenues of a firm-wage pair are increasing in output price and productivity, an increase in price reduces the threshold productivity for survival. With structural reforms (such as trade reforms) or even their reversal, the domestic relative price of one sector goes down and that of the other goes up. In the sector where the relative price goes up, the threshold productivity level falls and, as a result, the job creation rate rises and the job destruction rate falls. In the other sector, where the relative price goes down, the threshold productivity rises and, as a result, the job destruction rate rises and the job creation rate falls. Thus starting from a steady state of zero net job creation or destruction in each sector, we get net job creation in the sector with the price rise and net job destruction in the other sector. Since job creation takes time and job destruction can be instantaneous, the impact effect of a structural reform can be an increase in economy-wide unemployment.

## b. Endogenous structural reforms

In this study, we are looking at the impact of structural reforms on the labor-market outcomes of interest. However, whether structural reforms take place or not and their timing depend on economic and political factors and interactions between such factors. Macroeconomic conditions themselves might determine structural reforms. For example, poor macroeconomic performance and conditions (that would include high unemployment rates or alternatively high inflation rates) might lead governments to seek technical help from multilateral institutions, including the International Monetary Fund. Such technical assistance often comes with the conditionality of deep economic reforms, especially structural reforms. In certain other cases, high levels of government spending lead to high aggregate economic activity and rate of economic growth and a low unemployment rate. However, this high level of spending is possible through big budget deficits. Over time, national debt keeps building and at a certain point in time the size of the debt can reach crisis-like proportions. Such a debt crisis can lead the government to seek financial assistance from the International Monetary Fund, which comes attached with the strong conditionality of reforms. Reforms that were not politically viable earlier become politically viable. The external debt-to-GDP ratio can be a good instrument for reforms in our regressions because, while it triggers reforms, it does not, by itself, affect unemployment. It is government spending (and the budget deficit) that is related to unemployment and other labor-market outcomes. Note that while the budget deficit and government spending are flow variables, a country's overall and external debt are stock variables. In other words, the deficit can be high but the debt need not be high unless the high deficit is run for many years.

Related to the issue of the International Monetary Fund's role in reforms is the issue of "status-quo bias" that prevents reforms, which is very elegantly demonstrated by Fernandez and Rodrik (1991). This can happen even when the movers (all those who move from the import-competing sector to the export sector after the reform) and those who were in the export sector prior to the reform (and continue there after the reform) benefit and form the majority of voters in the presence of uncertainty about who ends up moving. In the words of Fernandez and Rodrik (1991), there is "individual-specific uncertainty" regarding who ends up moving and who ends up staying in the import-competing sector upon reforms. Let us say all those who are in the export sector prior to reforms gain from reforms. Let us assume 40 per cent of the population is in that sector to begin with. After reforms, this sector will utilize 70 per cent of the population. Each mover will gain $x$ and each person stuck in the import-competing sector will lose $y$. Let us assume $y > x$, in which case prior to reform, any producer in the import-competing sector initially views her expected change in welfare as $0.5(x - y) < 0$. Thus, we get a vote against the reform ex ante even if ex post a majority of the people benefit. When there is a debt crisis, which can be mitigated through financial assistance from the International Monetary Fund, the government might care more about obtaining this assistance than about popular support. Alternatively, the popular support could be affected by the possibility of this assistance in a crisis situation. Suppose this assistance conditional on structural reforms adds an amount $\Delta$ to the above expected change in welfare for every individual when a crisis is looming and can be prevented by International Monetary Fund assistance. If $\Delta$ is large enough (which could be the case if the possible crisis is severe and the potential International Monetary Fund assistance is substantial), then we can have $0.5(x - y) < 0 < 0.5(x - y) + \Delta$. Under these conditions, reforms will take place. A subsequent vote on whether reforms should continue or be reversed, even in the absence of further International Monetary Fund pressure or financial assistance, will support the continuation of reforms since the reforms have now already revealed the identities of the beneficiaries of these reform, i.e., the "individual-specific uncertainty" has been resolved.

A few macroeconomic models of political economy explain the delay in fiscal stabilization. Two such models are Alesina and Drazen (1991) and Drazen and Grilli (1993). In those models, reforms are like a public good. However, there are also costs that need to be incurred while implementing reforms. Just like any costly public good that generates benefits for everyone, reforms are faced with a free-rider problem. Everyone wants everybody else but themselves to incur the costs. In Alesina and Drazen (1991) and Drazen and Grilli (1993), there are two groups, both of which would benefit from a fiscal reform. In the absence of a fiscal reform, there are distortionary taxes that are costly and cannot raise enough revenues to close the fiscal deficit. A fiscal reform will raise direct taxes that are less distortionary and will improve and stabilize the macroeconomic climate. There is a war of attrition between the two groups in that each wants to out-wait the other one in the hopes that they will give in and incur most of the costs (will agree to bear a bigger share of the tax burden). The model is based on incomplete information available to a group about the foregone benefits of the other group. Thus, there is a delay until one group gives in. The costs rise over time due to a buildup of the fiscal debt. It is important to note here that any financial assistance from an international institution should be able to speed up stabilization in that it could increase the current benefit from reform for both groups.

Let us reiterate that the above models mean that the country's debt can be an important variable determining the likelihood and timing of reforms. These models also mean that if an economy is reformed, the likelihood of reversal is low. Therefore, whether the economy was in a state of reform or not in the previous period will determine the current state. The lagged value of the variable, which indicates whether an economy is reformed or not, is a possible instrument but could be a questionable one if the effect of reforms on labor-market outcomes happens with a lag. On the other hand, if labor-market outcomes change in anticipation of reforms, then this could be a good instrument.

Another factor that could potentially affect the likelihood of reforms is a country's terms of trade. The question that might exist about such an instrument is that terms of trade by themselves may be an important determinant of labor market outcomes, affecting the domestic labor-market outcomes directly and independently of the reforms taking place, not just through their effect on the possibility and the timing of reforms. For example, an adverse terms of trade shock increases real product wages because it creates a wedge between consumer prices, which is the relevant deflator for workers' wage indexation, and the value added price deflator, which is the relevant deflator for the employment decision. In other words, wages increase in excess of their full employment equilibrium and creates unemployment, wage hikes, and increase labor force participation.

What might break this direct relationship between unemployment and the terms of trade is the fact that prior to reforms, most developing countries were so closed that the external terms of trade that they faced would not have affected their labor-market outcomes in their state of virtual autarky. In addition, their trade policies (and other structural policies) tried to completely insulate them from fluctuations in their external terms of trade. Given that in our empirical analysis we will have a substantial number of observations pre- and post-reforms, this direct correlation between unemployment and terms of trade after controlling for reforms will break down.

We summarize here two models that show how changes in the terms of trade can affect the likelihood of reforms. Krishna and Mitra (2008) use a median-voter political-economy model to study this relationship. Consider an economy with two sectors, an import-competing sector and an exportable sector. Every individual in this economy is assumed to have his/her personal comparative advantage (relative productivity) in the production of these goods. In this model when the relative world price of the importable goes down (the relative world price of the exportable improves), amounting to a terms-of-trade improvement, the proportion of voters supporting the reform goes up. With a higher world relative price of the exportable, the relative payoff to working in the export sector goes up for everyone.

In another paper, Krishna and Mitra (2005) extend a Grossman-Helpman "Protection for Sale" lobbying model to include endogenous lobby formation by specific factor owners in the export sector in the presence of a pre-existing import-competing lobby. The authors of the paper show that an increase in the world price of the exportable increases the incentive for lobby formation by the producers of the export good, which neutralizes the import-competing lobby to bring about reform.

Another paper that arrives at the same result as above is an older paper by Coates and Ludema (2001). While they take a more black-box approach to lobbying than Grossman and Helpman (1994), they introduce some other "real-world complications" that are quite relevant in this context. In particular, they introduce negotiations between trading partners and ratification by each country's legislature.

We next move to the role of democratization in explaining liberalization. Milner and Kubota (2005) argue that democratization makes it harder for protectionist governments to maintain political support in developing countries (labor-abundant countries). Suppose individuals possess capital and labor in a two-sector, two-factor economy. In a labor-abundant country, freer trade through structural reform increases the reward to labor while reduces the return on capital. In a labor-abundant country where only the elite (the relative capital-rich individuals) can vote, we will see protectionist policies in place since those policies will benefit the scarce factor, capital, at the expense of the abundant factor, namely labor. Democratization gives the relatively capital-poor (people with fewer assets) the right to vote, which moves the country to more liberalized policies as it benefits the abundant factor, labor. Milner and Kubota (2005) find strong support for their hypothesis using data from developing countries for the period 1970-99. Also, Mitra, Thomakos and Ulubasoglu (2001) find that the weight the government puts on aggregate welfare relative to political contributions in a Grossman-Helpman "Protection for Sale" model has been higher with democratic governments than with dictatorships in Turkey. In the presence of lobbies that predominantly represent import-competing sectors, this result supports the positive relationship between democratization and liberalization. Thus a change in a country's democracy score over the last several years can be used as an instrument for structural reforms.

## 3. Econometric Methodology

We first run the following regression for each of our labor-market outcomes:

$$Y_{it} = \beta_0 + \beta_1 X_{it} + \sum_{j=-10}^{-1} \gamma_j PreDummy(j)_{it} + \sum_{j=1}^{10} \gamma_j PostDummy(j)_{it} + v_i + \varepsilon_{it} \quad (1)$$

where $i$ denotes the country, $t$ denotes the year, $Y$ denotes the labor-market outcome variable under consideration (unemployment rate, employment, female employment, wage rate, labor force participation rate or female labor force participation rate), $X$ includes control variables if specified (these regressions are run with and without controls), *PreDummy*(*j*) is a dummy variable equal to 1 *j* years prior to the reform (0 otherwise), *PostDummy*(*j*) is a dummy variable equal to 1 *j* years after the reform (0 otherwise), $v$ is the country fixed effect and $\varepsilon$ is the idiosyncratic error. The coefficient of each *PreDummy*($j$) and *PostDummy*($j$) represents the average level of the labor outcome variable in its corresponding time period relative to the average level the year of reform. The controls include real GDP, labor force participation rate, and working age population for all labor-market outcome variables except when the overall and female labor force participation rates are themselves the dependent variables. When the dependent variable is the overall of female labor force participation rate, the controls used are the real GDP and working age population. After running the above regressions, we present graphs that plot the estimated dummy variable regression coefficients against time where time is equal to $-1$ the year prior to the reform, 0 the year of reform, 1 the year after reform, etc.

We next run the following regression for each labor-market outcome variable:

$$Y_{it} = \alpha_0 + \alpha_1 POST_{it} + \alpha_2 time_{it} + \alpha_3 (time_{it})^2 + \alpha_4 X_{it} + u_i + \omega_{it} \qquad (2)$$

where, just as in the case of (1), $i$ denotes the country, $t$ denotes the year, $Y$ denotes the labor-market outcome variable under consideration (unemployment rate, employment, female employment, wage rate, labor force participation rate or female labor force participation rate), $X$ includes control variables if specified (these regressions are run with and without controls), $u$ is the country fixed effect and $\omega$ is the idiosyncratic error. *POST* is a dummy variable that takes the value 1 every year after the reform and 0 otherwise. Once again, the controls include real GDP, labor force participation rate, and working age population for all labor-market outcome variables except the overall and female labor force participation rates. When the dependent variable is the overall or female labor force participation rate, the controls used are the real GDP and working age population. Specification (2) above is run as a fixed effects, ordinary least squares regression with and without controls.

As argued in our theory section above, the implementation and timing of reforms can be endogenous to economic variables, especially to macroeconomic policies and conditions. Thus, our variable POST can be endogenous to the unemployment rate and other labor-market outcome variables of interest. To tackle this endogeneity problem, we run the same fixed-effects specification for each labor-market outcome variable, both with and without controls, as an instrumental variables (IV) regression using alternative sets of instruments. The variable that is instrumented is POST. POST is instrumented with external debt to GDP ratio, terms of trade, and the 5-year change in a country's democracy score. Alternatively, POST is instrumented with change in the world interest rate (proxied by the change in the US Treasury bill rate) interacted with the external debt to GDP ratio, terms of trade, and the 5-year change in a country's democracy score. Additionally we also include one-year lagged POST in our list of instruments. We have explained in the theory section, using economic intuition based on existing models, why these are potentially good instruments for POST.

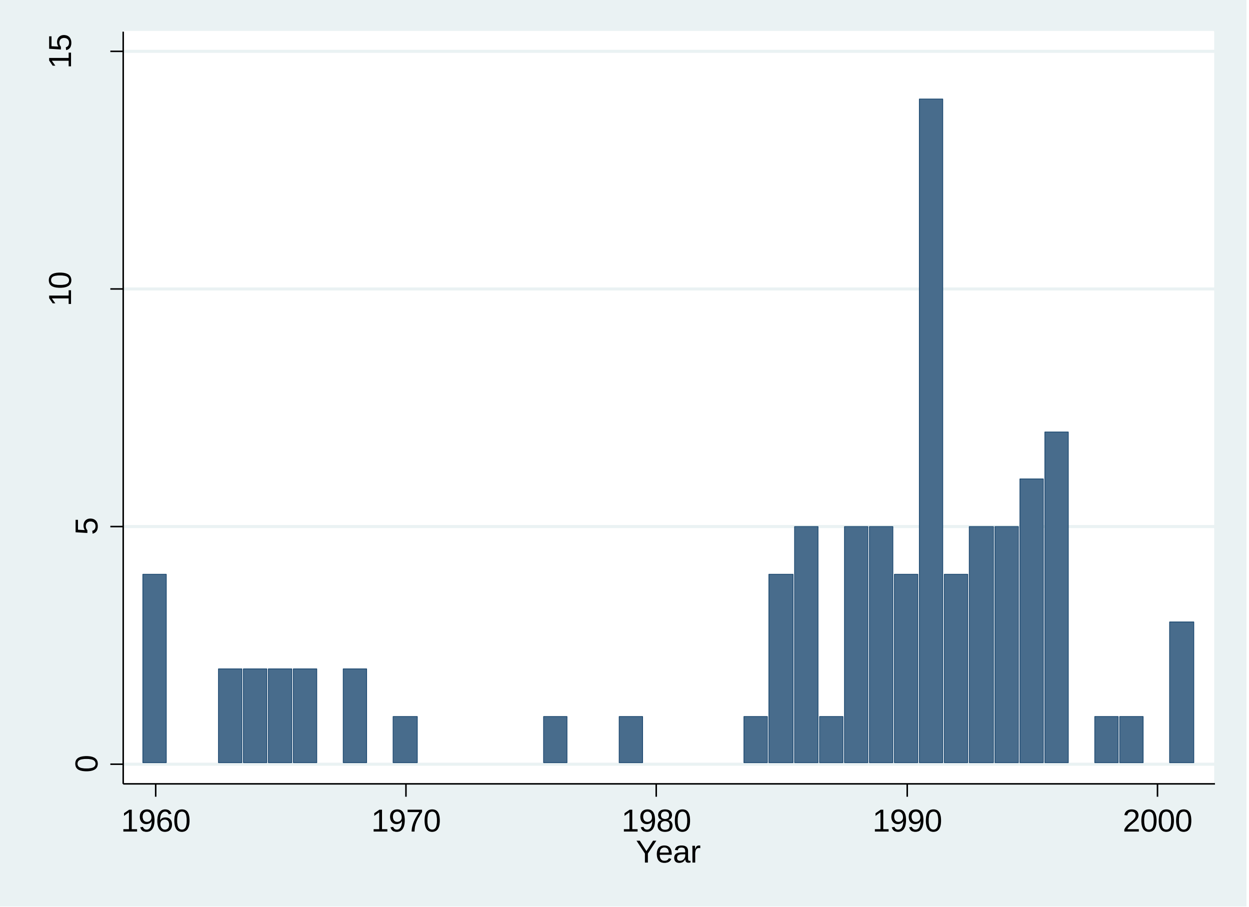
We also conduct econometric tests for the validity of these instruments. The first requirement is that these instruments should be fairly correlated with the endogenous variable we are trying to instrument. After partialling out the effects of the exogenous variables that are included in the second stage, a substantial proportion of the variation in the instrumented endogenous variable should be explained by its instruments. To ascertain that we calculate Shea's partial $R^2$. We also report the Kleibergen Paap F-statistic to test for weak instruments and compare them to critical values tabulated by Stock and Yogo (2005). Instruments are considered weak if the statistic exceeds the critical value. The second requirement is that each of the instruments is uncorrelated with the error term. To be able to verify that, we need our main equation of interest to be overidentified, that is, the number of instruments to exceed the number of right-hand side variables instrumented. If the equation is overidentified, then the Hansen J-statistic for overidentifying restrictions can be calculated (when not assuming i.i.d. errors). The null hypothesis of this test is that all the extra instruments are jointly exogenous (each is uncorrelated with the error term). If the p-value corresponding to the value of this test statistic exceeds 5 percent, the null hypothesis of joint exogeneity of these instruments cannot be rejected at the 5 percent significance level (and so on).

# 4. Data

The country sample and liberalization dates follow from Wacziarg and Welch (2008). Wacziarg and Welch (2008) revise and update Sachs and Warner's (1995) dates of liberalization through 2001. In principle, the liberalization date is the date after which all of the Sachs-Warner openness criteria are continuously met (however data limitations often imposed reliance on country case studies of trade policy). A country is classified as open if it displays none of the following characteristics: (1) average tariff rates of 40 percent of more; (2) nontariff barriers covering 40 percent or more of trade; (3) a black market exchange rate at least 20 percent lower than the official exchange rate; (4) a state monopoly on major exports; and (5) a socialist economic system.

Our structural reform index, *POST*, is a dummy variable that takes the value 1 every year after the date of liberalization and 0 otherwise. We call it a structural reform index because it proxies for the date when countries reached a threshold of broad reforms including macroeconomic stabilization, privatization, trade opening, as well as the end of interventionist states such as communism. For countries with multiple attempts at liberalization, the year of the final liberalization episode was used. Countries that liberalized their policies prior to 1960 are classified as "always open". Countries that had not liberalized by 2001 are classified as "always closed". Figure 1 presents a histogram of the number of countries liberalizing in each year. See appendix a for a complete list of the countries included in Wacziarg and Welch's (2008) dataset and the year of liberalization.

**Figure 1: Histogram of Year of Reform**




Notes: Figure 1 plots the number of countries that reformed each year from 1960 to 2001.

Six labor-market outcomes are considered separately as dependent variables of interest. The unemployment rate (UnempRate) is defined as the percentage of the labor force that is without work but available for and seeking employment. The series was constructed using data from the International Monetary Fund's International Financial Statistics (International Monetary Fund 2010), the World Bank's World Development Indicators (World Bank 2012) as well as the International Labor Organization's Key Indicators of the Labor Market (International Labor Organization 2012), the Organization for Economic Co-operation and Development's Labor Force Statistics (Organization for Economic Co-operation and Development 2012), and other regional agencies and country-specific sources. Employment (Emp) measured in millions, accessed from The Conference Board's Total Economy Database (The Conference Board 2012), includes employees, the self-employed, unpaid family members that are economically engaged, apprentices, and the military. Employment series for countries not available from The Conference Board's Total Economy Database were accessed from the World Bank's World Development Indicators as total employment aged 15 and older. Female Employment (FemaleEmp) measured in millions, accessed from the World Bank's World Development Indicators, is total female employment aged 15 and older. The wage index (WageIndex), accessed from the International Monetary Fund's International Financial Statistics, is an index of wage earnings with 2005 as the base year equal to 100.[17,18]

The labor force participation rate (LFPRate), accessed from the World Bank's World Development Indicators, is the labor force as a percentage of the working age population. Total labor force comprises people ages 15 and older who meet the International Labour Organization definition of the economically active population: all people who supply labor for the production of goods and services during a specified period. It includes both the employed and the unemployed. While national practices vary in the treatment of such groups as the armed forces and seasonal or part-time workers, in general the labor force includes the armed forces, the unemployed, and first-time job-seekers, but excludes homemakers and other unpaid caregivers and workers in the informal sector. The latter—the exclusion of informal workers from the definition of the labor force—turns out to be important for the interpretation of the econometric evidence discussed in the results below. Labor force participation rate series for countries not available from the World Bank's World Development Indicators were accessed from The International Monetary Fund's International Financial Statistics as the labor force as a percentage of the population aged 15 and older. The female labor force participation rate (FemaleLFPRate), accessed from the World Bank's World Development

[17] Ideally we would use real wages as a labor market indicator rather than nominal wages as captured by the wage index, since nominal wages are influenced by inflation responses of trade reforms. Unfortunately data on consumer prices is not available for our large sample of countries over the considered time period. In addition, one may wish to use real wages relative to productivity as a labor market indicator. However, as discussed above, overall productivity in the economy goes up after liberalization due to greater competition in international markets and lower productivity firms exiting, which leads to higher wages. Thus productivity is one channel through which reforms impact wages, and we do not wish to dampen this effect.

[18] Ideally we would also use wage data for unskilled versus skilled workers, as the Heckscher-Ohlin framework predicts that unskilled wages increase in response to reform in countries with abundance of unskilled labor, and considering wages for all workers may downward bias this effect. Unfortunately this data is not available for our large sample of countries over the considered time period.

Indicators, is the female labor force as a percentage of the female working age population.

Additional control variables are included. Real GDP (RealGDP) is measured in thousands of constant 2005 International Dollars and is accessed from the Penn World Tables version 7.0 (Heston et al. 2011) as PPP converted GDP per capital (chain series) constant 2005 prices multiplied by the population. Working age population (WorkingAgePop) measures the number of people who could potentially be economically active as the total population between the ages 15 to 64, accessed from the World Bank's World Development Indicators.

Four instrumental variables were considered to instrument $POST$. External debt to GDP ratio (IV1B_Debt) is total external debt (debt owed to nonresidents repayable in foreign currency, goods or services)[19] from the World Bank's World Development Indicators as a share of GDP. External debt to GDP ratio interacted with the change in world interest rates (proxied by the US Treasury bill rate) (IV1A_Debt), is the same as above only interacted with the one year change in the market yield on U.S. Treasury securities at 10-year constant maturity, available from the United States Federal Reserve. Terms of trade (IV2_TofT) is the net barter terms of trade index from the World Bank's World Development Indicators, calculated as the percentage ratio of the export unit value indexes to the import unit value indexes, measured relative to the base year 2000. The 5-year change in democracy score (IV4A_Dem) is calculated as the 5-year change in Polity IV's governing authority index (Gurr et al. 2012). Index scores range from -10 (most autocratic) to 10 (most democratic).

Table 1 reports summary statistics for each of the variables of interest.

[19] Total external debt is debt owed to nonresidents repayable in foreign currency, goods or services. Total external debt is the sum of public, publicly guaranteed, and private nonguaranteed long-term debt, use of International Monetary Fund credit and short-term debt. Short-term debt includes all debt having an original maturity of one year or less and interest in arrears on long-term debt. Data are in current U.S. dollars.

**Table 1: Summary Statistics**

| VARIABLES | N | max | min | mean | sd |
|---|---|---|---|---|---|
| UnempRate | 2,702 | 70.86 | 0.0410 | 8.537 | 6.487 |
| Emp | 5,150 | 758.5 | 0.0730 | 28.76 | 73.00 |
| LFPRate | 4,235 | 91.60 | 43.03 | 68.69 | 9.683 |
| FemaleLFPRate | 4,110 | 91.80 | 10.40 | 56.77 | 16.46 |
| WageIndex | 983 | 275.2 | 0 | 56.16 | 45.45 |
| FemaleEmp | 2,603 | 340.3 | 0.0290 | 7.199 | 29.01 |
| rGDP_PWT | 6,066 | 1.319e+10 | 170,251 | 2.483e+08 | 8.615e+08 |
| WorkingAgePop | 6,987 | 9.683e+08 | 111,527 | 2.053e+07 | 7.377e+07 |
| IV1A_Debt | 3,436 | 21.67 | -15.79 | -0.133 | 1.029 |
| IV1B_Debt | 3,436 | 18.23 | 0 | 0.661 | 0.920 |
| IV2_TofT | 3,009 | 721.1 | 21.28 | 111.6 | 41.19 |
| IV4A_Dem5 | 5,393 | 18 | -18 | 0.531 | 3.920 |

Notes: Table 1 presents, for each variable, the number of observations in the first column, the maximum in the second column, the minimum in the third column, the mean in the fourth column, and the standard deviation in the fifth column.

# 5. Results

## a. Time plots of labor-market outcomes

As previously mentioned, in Equation (1) the coefficient of each *PreDummy*($j$) and *PostDummy*($j$) represents the cross-country average of the level of the labor-market outcome variable in its corresponding time period relative to the level the year of reform where time is equal to $-1$ the year prior to the reform, 0 the year of reform, 1 the year after reform, etc. In Figure 2A, we specifically plot the cross-country average relative unemployment rate in each year. There are no controls in the estimation of (1), other than country fixed effects, in Panel A, while additional controls in Panel B include real GDP, labor force participation rate, and working age population. Both panels clearly show that, while on average unemployment fluctuates both before and after reforms, the post-reform average level of the unemployment rate looks distinctly higher than the pre-reform average unemployment rate. While the unemployment rate attains its lowest value a couple of years prior to reforms, it reaches its maximum level a couple of years after the reform, with the difference in the unemployment rates between these two points being 2.5 percentage points. This is just a documentation of empirical regularity rather than a statement of causation. If reforms accompanied by austerity measures are carried out to get out of a debt crisis arising in turn from many years of extravagant government spending (leading to high aggregate output and low unemployment), then clearly it is not reforms that are causing the higher unemployment. Also, a stock market crash can happen in anticipation of a bad real economic performance, i.e., the stock market is a good predictor of future economic performance. A poor financial situation and a poor performance of stocks often go together, while a financial crisis can trigger reforms. In this case again reforms cannot be responsible for higher unemployment rates. Despite these possibilities, we see from both Panel A (without controls other than fixed effects) and Panel B (with additional fixed effects) of Figure 2B that employment levels post-reform are way higher than those pre-reform. Also, after the reforms, employment has been rising consistently and relatively steeply for a long period of time. This is also true with the wage index presented in Figure 2C. Pre-reform levels are lower than post-reform levels by a wide margin. In addition, the post-reform wage index has a relatively much steeper and much more consistent upward trend.

We next look at the overall labor force participation rate in Figure 2D. In Panel A, we find a declining trend prior to the reform and the labor force participation rate exhibits an upward trend after the reform. In Panel B, while there are more fluctuations throughout, the trend on average is upward sloping after the reform, but was downward sloping prior to the reform. In Figure 2E, the female labor force participation rate also shows an upward trend with consistently positive levels relative to the reform year after the reform. Before the reform, negative relative female labor force participation rates are dominant in both Panel A and Panel B.

To dig a little deeper into the female labor-market outcomes, we plot the average female employment level each year relative to the reform year after controlling for country fixed effects in Panel A of Figure 2F and then in Panel B using the additional controls we have been using so far. With no controls other than fixed effects, female employment levels on average are higher before reforms in Panel A as compared to the pre-reform levels. In Panel B, while on average female employment level are higher pre-reform than post-reform, the negative trend pre-reform is replaced by a positive trend post-reform.

**Figure 2.A: Dependent Variable: Unemployment Rate**

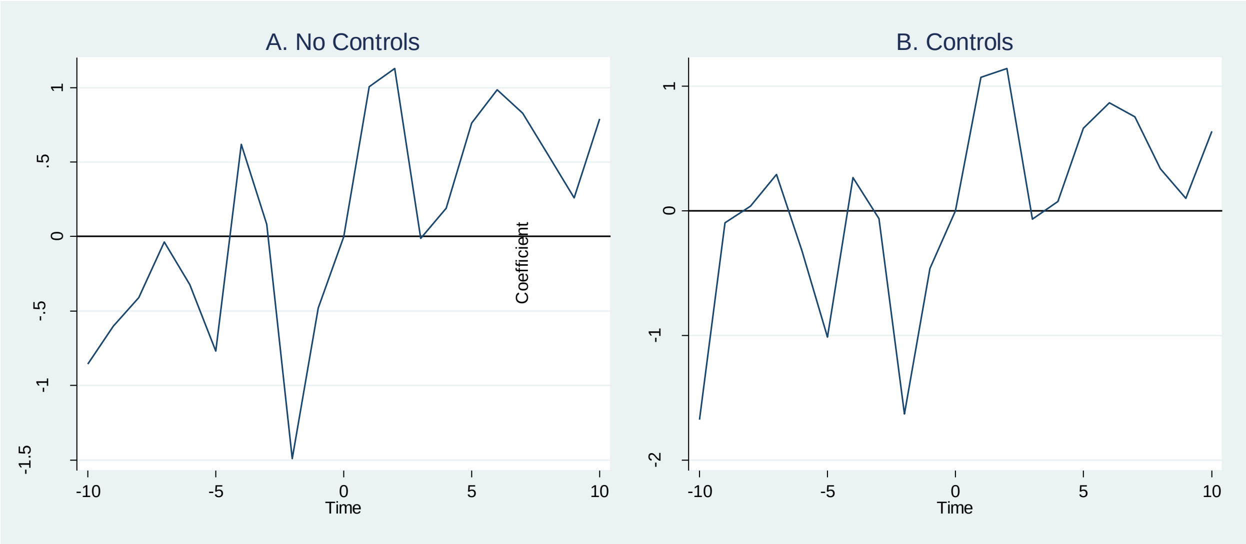



**Figure 2.B: Dependent Variable: Employment**

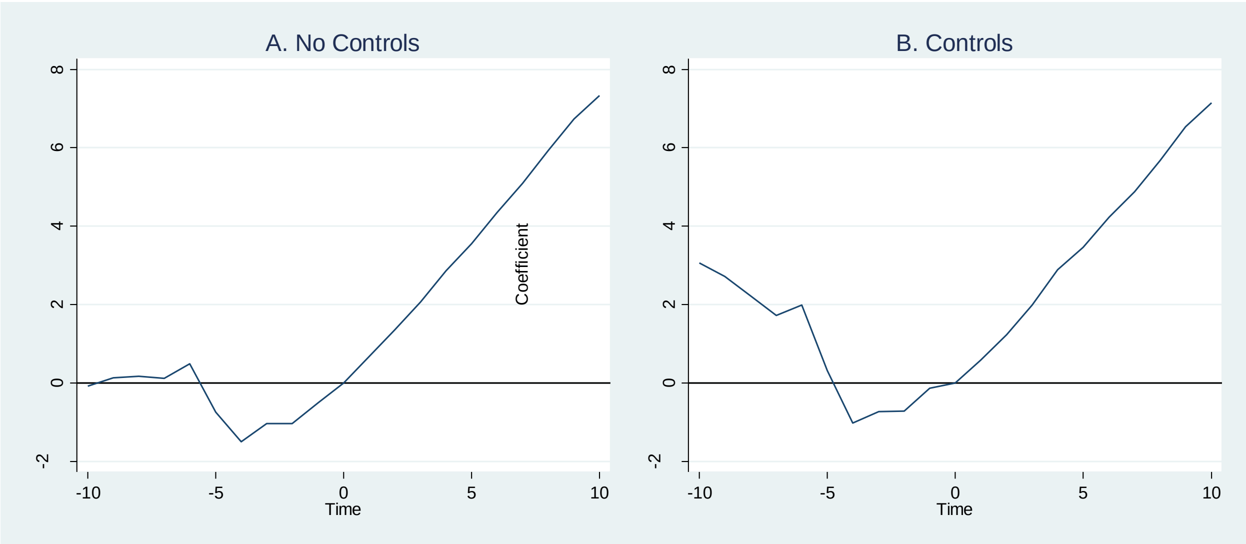


**Figure 2.C: Dependent Variable: Wage Index**

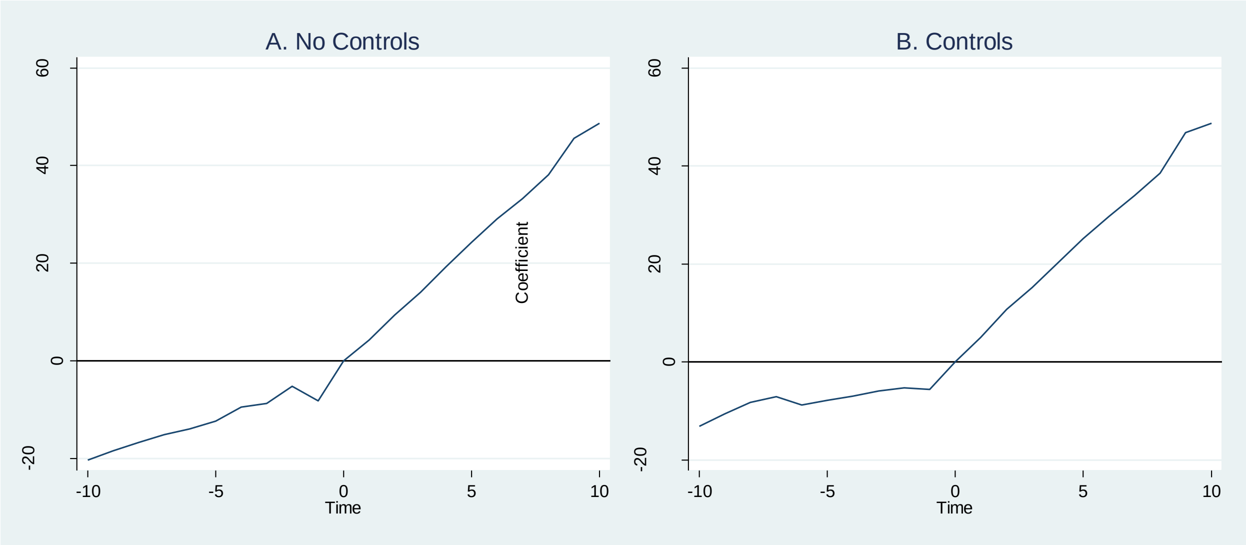



**Figure 2.D: Dependent Variable: Labor Force Participation Rate**

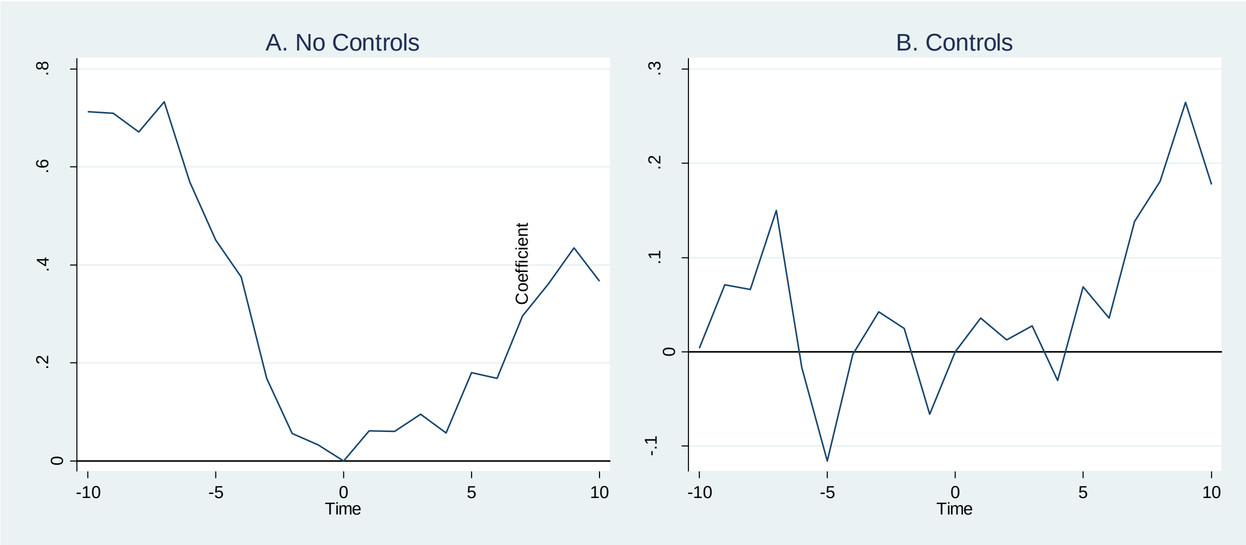



**Figure 2.E: Dependent Variable: Female Labor Force Participation Rate**

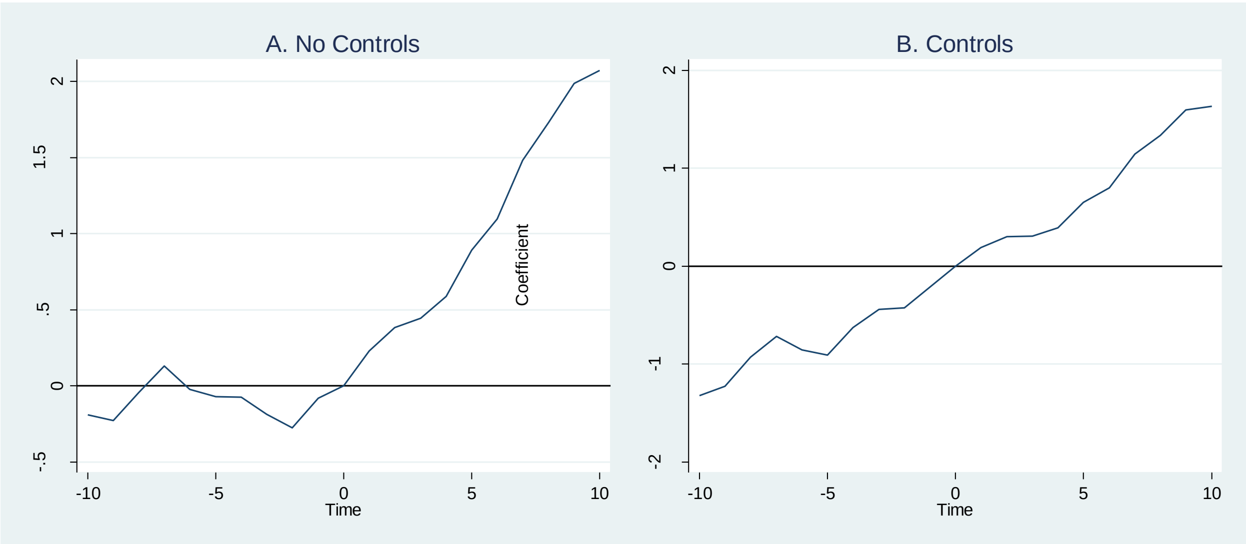
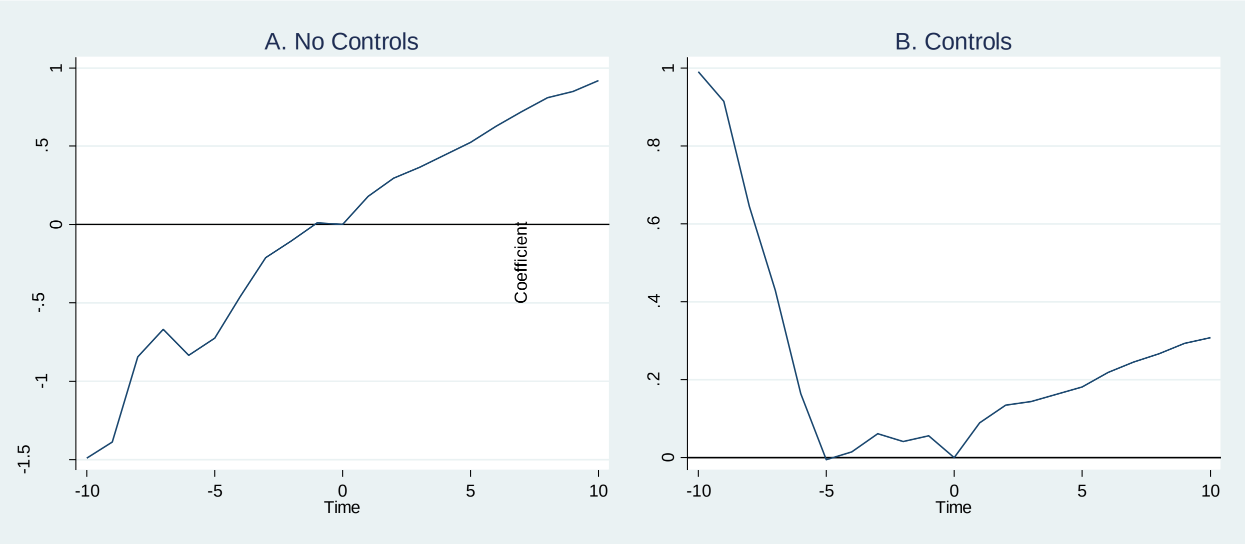


**Figure 2.F: Dependent Variable: Female Employment**




Notes: Figure 2 plots, for each dependent variable, the estimated dummy variable regression coefficients from the regression specification: $DependentVariable_{it} = \beta_0 + \beta_1 X_{it} + \sum_{j=-10}^{-1} \gamma_j PreDummy(j)_{it} + \sum_{j=1}^{10} \gamma_j PostDummy(j)_{it} + v_i + \varepsilon_{it}$ , where $i$ denotes the country, $t$ denotes the year, $X$ includes control variables if specified, $PreDummy(j)$ is a dummy variable equal to 1 $j$ years prior to the reform, $PostDummy(j)$ is a dummy variable equal to 1 $j$ years after the reform, $v$ is the fixed effect and $\varepsilon$ is the idiosyncratic error. Time is equal to $-1$ the year prior to the reform, 0 the year of reform, 1 the year after reform, etc. Each coefficient represents the average level of the dependent variable in its corresponding time period relative to the average level the year of reform. The controls in Panel B include real GDP and working age population in all regressions as well as labor force participation rate with the exception of the regression in which this appears as the dependent variable.

On average, after reforms, the unemployment rates are higher but all other labor-market outcomes on average look better both in terms of trends and levels. This does not mean anything in terms of causation. However, these are empirical regularities that are hard to ignore. Next we perform some OLS and IV fixed effects regressions to see if we can identify any causal effects of reforms on labor-market outcomes and, if so, whether this is qualitatively similar to the patterns observed in the plots presented in Figures 2A-2F.

## b. Regression results

Tables 2 through 7 present the results of the estimation of specification (2). In all these regressions we control for time and time squared. While five of the regressions (those presented in the odd-numbered columns) do not have any other controls besides the country fixed effects (FE), the remaining five regressions have other controls, namely working age population, the real GDP from the Penn World Tables and the labor force participation rate. The OLS-FE regression results are presented in columns (1) and (2).

The OLS-FE regression results presented in columns (1) and (2) of the tables show similar outcomes to what we have seen in our figures above. In Table 2, the dependent variable is the unemployment rate. The coefficient of POST is positive and significant and shows that the unemployment rate on average is about a percentage point higher after the reforms than before. This is true even though we are controlling for a time trend and the square of time, both of which are statistically insignificant. This shows there is clearly no time pattern of the unemployment rate except that the average level post-reform is clearly higher than the average level pre-reform. An expansion of the labor force participation rate is accompanied by a reduction in the unemployment rate. For the OLS regressions, when the observations are restricted to only those for which there is information on all the basic and control variables and all the IVs, the OLS sample gets much more restricted. POST is now highly insignificant but is also negative. These results are not presented in our tables but are available upon request.

When employment (Table 3), female employment (Table 7) and the wage index (Table 4) are the dependent variables, the OLS-FE regression results show that the coefficient of POST is positive and significant when the additional control variables are included in the regression (while for employment the coefficient of POST is also positive and significant without the additional control variables). Employment (total and female) and wages are higher on average in economies following structural reforms. Restricting the sample for the OLS regressions to the sample used in the IV regressions for employment and the wage index makes the results even stronger (not shown in our tables). However, the OLS-FE results for the labor force participation rates (total and female) are weak.

To tackle the endogeneity of POST discussed in sections 2 and 3, we run the same fixed-effects specifications for each labor-market outcome variable as IV regressions, using alternative sets of instruments. The variable that is instrumented is POST. The instruments for regressions presented in columns (3) and (4) are external debt to GDP ratio, terms of trade, and the 5-year change in democracy score. The one-year lagged value of POST is added as an instrument in columns (5) and (6). Columns (7), (8), (9) and (10) are respectively the same as (3), (4), (5) and (6), except that now the external debt to GDP ratio as an instrument is replaced with the interaction between the change in the world interest rate (proxied by the US Treasury bill rate) and external debt over GDP.

A valid instrument is one that is correlated with the endogenous regressor yet orthogonal to the errors. To determine the quality of the instruments, three empirical tests are used that provide evidence of the instruments' validity. First, to assess the correlation of the instruments with the endogenous regressor, it is sufficient to examine the significance of the excluded instruments in the first stage regression. A commonly used statistic is the $R^2$ of the first-stage regression referred to as the partial-$R^2$ (Shea 1997). Although there is no threshold level, the instruments should be sufficiently relevant to explain a significant share of the variance of the endogenous regressor. As a rule of thumb, an estimated equation that yields a partial-$R^2$ lower than 10 percent indicates a "weak instruments" problem. Second, we also report the Kleibergen Paap F-statistic to test for weak instruments and compare them to critical values tabulated by Stock and Yogo (2005). Instruments are considered weak if the statistic exceeds the critical value.

Third, whether the instruments are orthogonal to the errors can be tested in an overidentified model where the number of instruments is greater than the number of endogenous regressors using the Hansen J-statistic. The Hansen J-statistic is from a test of the hypothesis that the instruments are uncorrelated with the error term or that the overidentification restrictions are valid. A rejection would call this hypothesis into question. The partial-$R^2$, the Kleibergen Paap F-statistic, and the Hansen J p-value are reported in each of the IV results.

When the unemployment rate is the dependent variable, POST consistently has a negative coefficient sign, but is only significant in three of the eight regressions. Only in columns (5), (6), (9) and (10), where lagged POST is also included as an instrument, do both the first-stage partial $R^2$, the Kleibergen-Paap F statistic, and the Hansen J p-value indicate that the set of instruments possesses both ideal characteristics, namely a fair degree of correlation with the instrumented variable and the joint orthogonality of the instruments with respect to the error term. (The first stage regression results are available in appendix b.) The IV results from these columns imply that reforms on average lower the unemployment rate by about 1.5 percentage points, even when controlling for the labor force participation rate of the formal sector.

We next move to the employment level as a dependent variable (Table 3). The right-hand side variables and the exact set of instruments used in each column are the same as in the corresponding column numbers of Table 2. In most columns, the coefficient of POST is positive and quite significant in four of the eight IV-FE columns. This is despite the fact that, in many of these columns, time or time squared has a positive and highly statistically significant coefficient. This makes the result especially strong as these controls cover demographical changes and business cycle effects on labor markets. There is also strong evidence for the positive effect of the size of the working age population. In columns (5), (6), (9) and (10), the first-stage partial $R^2$, the Kleibergen-Paap F statistic and the Hansen J p-value indicate that the set of instruments exhibits both ideal characteristics required of them, namely a fair degree of correlation with the instrumented variable and the joint orthogonality of the instruments to the error term. In columns (9) and (10), where the set of IVs consists of the external debt to GDP ratio interacted with the US Treasury bill rate (to proxy for the world interest rate), terms of trade, the 5-year change in democracy score and lagged POST, the results of the Hansen J p-value are even stronger. Overall, the results are quite robust in those four columns, showing that structural reforms lead to an average employment increase of 4 million for the entire decade after the reforms relative to the decade before the reforms. In addition, because the labor force participation rate of the formal sector is being controlled for, the less conclusive results on unemployment after reforms suggests that part of the increase in employment may be from informal workers. The sample is heavily weighted by developing countries where informality is ubiquitous (since most developed countries were reformed throughout the period).

The wage index is our next dependent variable (Table 4). We follow the same sequence of columns in terms of the methods of estimation, right-hand side variables and IVs used. Across all columns in Table 4, the coefficient of POST is positive and is quite significant in five of the eight IV-FE columns. Once again, this is despite the fact that in most of these columns, both time and time squared have positive and highly statistically significant coefficients. These are very strong results. There is also strong evidence in the IV regression for the negative effect of the size of the working age population, while the sign of the real GDP from the Penn World Tables is what one would expect and is significant. Everywhere the first-stage partial $R^2$, Kleibergen-Paap F statistic, and the Hansen J p-value indicate that the set of instruments exhibits the ideal characteristics required of them (although in columns (4) and (5) the Kleibergen-Paap F statistic is only significant at the 20 to 25 percent level, depending on whether the maximal relative bias or maximal size test is used). In the first stages of column (3), the external debt-to-GDP ratio and terms of trade significantly increase the probability of reform. The lagged POST variable is also shown to significantly increase the probability of reform for its respective regressions.

In Tables 5-6, we look at the overall and female labor force participation rate of formal workers. While there seems to be a positive time trend throughout, POST is insignificant throughout. In half of the IV columns, the three test results indicate that the requirements for IVs are satisfied. In Table 7, we dig deeper into female labor-market outcomes by looking at their employment levels. Only cases where POST is significant when instrumented are those where the instruments exhibit both ideal characteristics required of them, although the null of joint exogeneity is not reject in any of these columns by the Hansen J-statistic. The coefficients however are not strong and, in our opinion, quite inconclusive. Overall, based on these regression results and Figure 2F discussed above, we cannot rule out the positive effect of structural reforms on female employment as well.

The evidence suggests that structural reforms were associated with a simultaneous reduction in unemployment, as well as higher wages and higher labor force participation of formal workers. While this increase in wages incentivizes workers to join the labor market, from a firm's perspective, higher real wages leads to higher unemployment unless there are compensation productivity advances. Thus these results may evidence firm-level increases in productivity that allows them to pay higher wages while employing more workers. It may also signal the entry of new firms or the expansion of existing firms that now have access to a larger external market.

In sum, higher economy-wide real wages after structural reforms improve job prospects and provide incentives for workers to find jobs, consistent with higher levels of employment after reforms. But controlling for the labor force participation of formal workers and the weak results on unemployment suggest this increase in employment could be coming from informal workers entering into the labor force.

**Table 2: Unemployment Rate**

| VARIABLES | (1)<br>OLS-FE | (2)<br>OLS-FE | (3)<br>IV-FE | (4)<br>IV-FE | (5)<br>IV-FE | (6)<br>IV-FE | (7)<br>IV-FE | (8)<br>IV-FE | (9)<br>IV-FE | (10)<br>IV-FE |
|---|---|---|---|---|---|---|---|---|---|---|
| POST | 1.00* | 1.12* | -6.52** | -5.11 | -1.54* | -1.56* | -7.30 | -5.17 | -1.40 | -1.42 |
| | (0.52) | (0.60) | (3.03) | (3.13) | (0.90) | (0.92) | (6.78) | (5.27) | (0.91) | (0.93) |
| Time | 0.00 | -0.02 | 0.49** | 0.34 | 0.12* | 0.07 | 0.55 | 0.34 | 0.11 | 0.06 |
| | (0.10) | (0.11) | (0.24) | (0.26) | (0.07) | (0.08) | (0.50) | (0.39) | (0.07) | (0.08) |
| Timesq | -0.00 | -0.00 | -0.01 | -0.01 | -0.01 | -0.01 | -0.01 | -0.01 | -0.01 | -0.01 |
| | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) |
| WorkingAgePop | | -0.00 | | -0.00 | | -0.00 | | -0.00 | | -0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| rGDP_PWT | | 0.00 | | 0.00 | | 0.00 | | 0.00 | | 0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| LFPRate | | -0.29* | | 0.06 | | 0.07 | | 0.06 | | 0.07 |
| | | (0.16) | | (0.17) | | (0.17) | | (0.17) | | (0.17) |
| Observations | 849 | 765 | 514 | 514 | 514 | 514 | 514 | 514 | 514 | 514 |
| R-squared | 0.01 | 0.03 | -0.13 | -0.07 | 0.00 | 0.01 | -0.17 | -0.07 | 0.00 | 0.01 |
| Number of Countries | 67 | 59 | 48 | 48 | 48 | 48 | 48 | 48 | 48 | 48 |
| Partial R-Squared | | | 0.03 | 0.03 | 0.42 | 0.42 | 0.02 | 0.02 | 0.42 | 0.42 |
| Hansen J p-value | | | 0.86 | 0.65 | 0.49 | 0.58 | 0.43 | 0.43 | 0.58 | 0.54 |
| Kleibergen-Paap F statistic | | | 5.45 | 5.42 | 500.47 | 466.94 | 3.07 | 3.18 | 381.01 | 368.91 |

Notes: Table 2 presents the estimated OLS and second-stage instrumental variables regression coefficients for unemployment rate as the dependent variable for different specifications. POST is instrumented with external debt to GDP ratio, terms of trade, and 5-year change in democracy score in columns (3) and (4). In columns (5) and (6) the instruments also include lagged POST. POST is instrumented with external debt to GDP ratio interacted with the change in world interest rates (proxied by the US Treasury bill rate), terms of trade, and 5-year change in democracy score in columns (7) and (8). In columns (9) and (10) the instruments also include lagged POST. Standard errors clustered at the country level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

## Table 3: Employment

| VARIABLES | (1) OLS-FE | (2) OLS-FE | (3) IV-FE | (4) IV-FE | (5) IV-FE | (6) IV-FE | (7) IV-FE | (8) IV-FE | (9) IV-FE | (10) IV-FE |
|---|---|---|---|---|---|---|---|---|---|---|
| POST | 0.99* | 1.31* | 13.62 | 8.44 | 4.27** | 4.18** | 2.75 | -2.02 | 4.25** | 4.16** |
| | (0.57) | (0.68) | (12.72) | (8.73) | (1.84) | (1.71) | (11.55) | (8.07) | (1.83) | (1.71) |
| Time | 0.31*** | 0.15** | -0.31 | -0.19 | 0.42*** | 0.14 | 0.54 | 0.63 | 0.42*** | 0.14 |
| | (0.07) | (0.07) | (0.94) | (0.69) | (0.12) | (0.13) | (0.87) | (0.67) | (0.13) | (0.13) |
| Timesq | 0.04*** | 0.05*** | 0.07** | 0.08** | 0.06*** | 0.07*** | 0.06* | 0.06** | 0.06*** | 0.07*** |
| | (0.01) | (0.02) | (0.04) | (0.03) | (0.02) | (0.02) | (0.03) | (0.03) | (0.02) | (0.02) |
| WorkingAgePop | | 0.00** | | 0.00* | | 0.00* | | 0.00* | | 0.00* |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| rGDP_PWT | | -0.00 | | -0.00 | | -0.00 | | -0.00 | | -0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| LFPRate | | 0.23 | | -0.16 | | -0.18 | | -0.21 | | -0.18 |
| | | (0.22) | | (0.56) | | (0.54) | | (0.56) | | (0.54) |
| Observations | 1,352 | 1,023 | 724 | 723 | 724 | 723 | 724 | 723 | 724 | 723 |
| R-squared | 0.17 | 0.33 | 0.20 | 0.35 | 0.31 | 0.38 | 0.31 | 0.36 | 0.31 | 0.38 |
| Number of Countries | 84 | 72 | 60 | 60 | 60 | 60 | 60 | 60 | 60 | 60 |
| Partial R-Squared | | | 0.02 | 0.02 | 0.41 | 0.41 | 0.02 | 0.02 | 0.41 | 0.41 |
| Hansen J p-value | | | 0.06 | 0.06 | 0.11 | 0.13 | 0.54 | 0.36 | 0.71 | 0.50 |
| Kleibergen-Paap F statistic | | | 2.89 | 2.88 | 518.30 | 513.40 | 3.41 | 3.52 | 509.16 | 504.79 |

Notes: Table 3 presents the estimated OLS and second-stage instrumental variables regression coefficients for employment as the dependent variable for different specifications. POST is instrumented with external debt to GDP ratio, terms of trade, and 5-year change in democracy score in columns (3) and (4). In columns (5) and (6) the instruments also include lagged POST. POST is instrumented with external debt to GDP ratio interacted with the change in world interest rates (proxied by the US Treasury bill rate), terms of trade, and 5-year change in democracy score in columns (7) and (8). In columns (9) and (10) the instruments also include lagged POST. Standard errors clustered at the country level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

## Table 4: Wage Index

| VARIABLES | (1) OLS-FE | (2) OLS-FE | (3) IV-FE | (4) IV-FE | (5) IV-FE | (6) IV-FE | (7) IV-FE | (8) IV-FE | (9) IV-FE | (10) IV-FE |
|---|---|---|---|---|---|---|---|---|---|---|
| POST | 3.46 | 4.75* | 72.44*** | 39.13** | 26.72 | 19.82 | 39.17** | 17.00** | 27.96* | 18.78 |
| | (2.89) | (2.71) | (13.24) | (18.09) | (17.91) | (15.35) | (15.54) | (8.20) | (14.91) | (12.21) |
| Time | 3.29*** | 2.81*** | -0.17 | 2.44** | 3.30*** | 3.60*** | 2.36 | 3.77*** | 3.21*** | 3.66*** |
| | (0.35) | (0.42) | (0.96) | (0.95) | (0.95) | (0.95) | (1.51) | (0.75) | (0.74) | (0.80) |
| Timesq | 0.15*** | 0.18*** | 0.33*** | 0.17** | 0.27*** | 0.15*** | 0.29*** | 0.14*** | 0.27*** | 0.15*** |
| | (0.05) | (0.04) | (0.11) | (0.07) | (0.06) | (0.05) | (0.04) | (0.05) | (0.06) | (0.05) |
| WorkingAgePop | | 0.00 | | -0.00** | | -0.00** | | -0.00** | | -0.00** |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| rGDP_PWT | | 0.00** | | 0.00*** | | 0.00*** | | 0.00*** | | 0.00*** |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| LFPRate | | 0.10 | | 2.41 | | 2.20 | | 2.17 | | 2.19 |
| | | (0.65) | | (2.13) | | (1.92) | | (1.88) | | (1.90) |
| Observations | 314 | 260 | 66 | 66 | 66 | 66 | 66 | 66 | 66 | 66 |
| R-squared | 0.76 | 0.87 | 0.65 | 0.86 | 0.84 | 0.90 | 0.81 | 0.90 | 0.84 | 0.90 |
| Number of Countries | 21 | 18 | 9 | 9 | 9 | 9 | 9 | 9 | 9 | 9 |
| Partial R-Squared | | | 0.18 | 0.19 | 0.47 | 0.44 | 0.16 | 0.10 | 0.49 | 0.47 |
| Hansen J p-value | | | 0.58 | 0.19 | 0.31 | 0.28 | 0.22 | 0.15 | 0.35 | 0.25 |
| Kleibergen-Paap F statistic | | | 116.92 | 7.78 | 11460.74 | 574.38 | 8.24 | 36.22 | 20331.37 | 33238.61 |

Notes: Table 4 presents the estimated OLS and second-stage instrumental variables regression coefficients for wage index as the dependent variable for different specifications. POST is instrumented with external debt to GDP ratio, terms of trade, and 5-year change in democracy score in columns (3) and (4). In columns (5) and (6) the instruments also include lagged POST. POST is instrumented with external debt to GDP ratio interacted with the change in world interest rates (proxied by the US Treasury bill rate), terms of trade, and 5-year change in democracy score in columns (7) and (8). In columns (9) and (10) the instruments also include lagged POST. Standard errors clustered at the country level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

## Table 5: Labor Force Participation Rate

| VARIABLES | (1) OLS-FE | (2) OLS-FE | (3) IV-FE | (4) IV-FE | (5) IV-FE | (6) IV-FE | (7) IV-FE | (8) IV-FE | (9) IV-FE | (10) IV-FE |
|---|---|---|---|---|---|---|---|---|---|---|
| POST | 0.04 | -0.01 | -0.44 | -0.31 | -0.24 | -0.25 | 0.05 | 0.29 | -0.24 | -0.24 |
| | (0.18) | (0.19) | (3.38) | (3.13) | (0.25) | (0.24) | (3.90) | (3.49) | (0.25) | (0.24) |
| Time | -0.03 | 0.01 | 0.12 | 0.11 | 0.11*** | 0.10*** | 0.09 | 0.06 | 0.11*** | 0.10*** |
| | (0.04) | (0.04) | (0.26) | (0.23) | (0.03) | (0.04) | (0.30) | (0.26) | (0.03) | (0.04) |
| Timesq | 0.01* | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| WorkingAgePop | | 0.00 | | 0.00 | | 0.00 | | 0.00 | | 0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| rGDP_PWT | | -0.00 | | 0.00 | | 0.00 | | 0.00 | | 0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| Observations | 1,430 | 1,333 | 960 | 960 | 960 | 960 | 960 | 960 | 960 | 960 |
| R-squared | 0.01 | 0.03 | 0.14 | 0.15 | 0.15 | 0.15 | 0.15 | 0.14 | 0.15 | 0.15 |
| Number of Countries | 73 | 73 | 62 | 62 | 62 | 62 | 62 | 62 | 62 | 62 |
| Partial R-Squared | | | 0.01 | 0.01 | 0.41 | 0.41 | 0.01 | 0.01 | 0.41 | 0.41 |
| Hansen J p-value | | | 0.39 | 0.37 | 0.50 | 0.43 | 0.43 | 0.38 | 0.56 | 0.50 |
| Kleibergen-Paap F statistic | | | 8.40 | 8.12 | 1945.14 | 1934.15 | 2.79 | 2.71 | 1904.93 | 1898.33 |

Notes: Table 5 presents the estimated OLS and second-stage instrumental variables regression coefficients for labor force participation rate as the dependent variable for different specifications. POST is instrumented with external debt to GDP ratio, terms of trade, and 5-year change in democracy score in columns (3) and (4). In columns (5) and (6) the instruments also include lagged POST. POST is instrumented with external debt to GDP ratio interacted with the change in world interest rates (proxied by the US Treasury bill rate), terms of trade, and 5-year change in democracy score in columns (7) and (8). In columns (9) and (10) the instruments also include lagged POST. Standard errors clustered at the country level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

## Table 6: Female Labor Force Participation Rate

| VARIABLES | (1) OLS-FE | (2) OLS-FE | (3) IV-FE | (4) IV-FE | (5) IV-FE | (6) IV-FE | (7) IV-FE | (8) IV-FE | (9) IV-FE | (10) IV-FE |
|---|---|---|---|---|---|---|---|---|---|---|
| POST | 0.04 | -0.02 | 0.10 | 1.32 | -0.34 | -0.35 | 1.00 | 2.76 | -0.34 | -0.35 |
| | (0.22) | (0.24) | (5.45) | (5.24) | (0.35) | (0.34) | (6.80) | (6.30) | (0.35) | (0.34) |
| Time | 0.11** | 0.15** | 0.24 | 0.13 | 0.28*** | 0.26*** | 0.18 | 0.03 | 0.28*** | 0.26*** |
| | (0.05) | (0.06) | (0.41) | (0.38) | (0.05) | (0.06) | (0.52) | (0.46) | (0.05) | (0.06) |
| Timesq | 0.01*** | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| WorkingAgePop | | 0.00 | | -0.00 | | -0.00 | | -0.00 | | -0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| rGDP_PWT | | 0.00 | | 0.00 | | 0.00 | | 0.00 | | 0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| Observations | 1,419 | 1,322 | 960 | 960 | 960 | 960 | 960 | 960 | 960 | 960 |
| R-squared | 0.07 | 0.19 | 0.34 | 0.33 | 0.34 | 0.35 | 0.32 | 0.27 | 0.34 | 0.35 |
| Number of Countries | 72 | 72 | 62 | 62 | 62 | 62 | 62 | 62 | 62 | 62 |
| Partial R-Squared | | | 0.01 | 0.01 | 0.41 | 0.41 | 0.01 | 0.01 | 0.41 | 0.41 |
| Hansen J p-value | | | 0.40 | 0.39 | 0.48 | 0.35 | 0.58 | 0.57 | 0.48 | 0.36 |
| Kleibergen-Paap F statistic | | | 8.40 | 8.12 | 1945.14 | 1934.15 | 2.79 | 2.71 | 1904.93 | 1898.33 |

Notes: Table 6 presents the estimated OLS and second-stage instrumental variables regression coefficients for female labor force participation rate as the dependent variable for different specifications. POST is instrumented with external debt to GDP ratio, terms of trade, and 5-year change in democracy score in columns (3) and (4). In columns (5) and (6) the instruments also include lagged POST. POST is instrumented with external debt to GDP ratio interacted with the change in world interest rates (proxied by the US Treasury bill rate), terms of trade, and 5-year change in democracy score in columns (7) and (8). In columns (9) and (10) the instruments also include lagged POST. Standard errors clustered at the country level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

## Table 7: Female Employment

| VARIABLES | (1)<br>OLS-FE | (2)<br>OLS-FE | (3)<br>IV-FE | (4)<br>IV-FE | (5)<br>IV-FE | (6)<br>IV-FE | (7)<br>IV-FE | (8)<br>IV-FE | (9)<br>IV-FE | (10)<br>IV-FE |
|---|---|---|---|---|---|---|---|---|---|---|
| POST | -0.01 | 0.18* | 0.76 | 0.76 | 0.12 | 0.57** | 4.27 | 0.63 | 0.12 | 0.57** |
| | (0.13) | (0.10) | (2.47) | (0.67) | (0.26) | (0.23) | (5.67) | (0.92) | (0.26) | (0.23) |
| Time | 0.11** | -0.02 | 0.08 | -0.09 | 0.15** | -0.07* | -0.29 | -0.08 | 0.15** | -0.07* |
| | (0.05) | (0.02) | (0.26) | (0.09) | (0.06) | (0.04) | (0.59) | (0.11) | (0.06) | (0.04) |
| Timesq | -0.00 | 0.00* | 0.00 | 0.01 | -0.00 | 0.01** | 0.02 | 0.01 | -0.00 | 0.01** |
| | (0.00) | (0.00) | (0.01) | (0.01) | (0.01) | (0.00) | (0.03) | (0.01) | (0.01) | (0.00) |
| WorkingAgePop | | 0.00*** | | 0.00*** | | 0.00*** | | 0.00*** | | 0.00*** |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| rGDP_PWT | | -0.00 | | 0.00 | | 0.00 | | 0.00 | | 0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| LFPRate | | 0.04*** | | 0.07* | | 0.07** | | 0.07* | | 0.07** |
| | | (0.01) | | (0.04) | | (0.04) | | (0.04) | | (0.04) |
| Observations | 800 | 785 | 575 | 575 | 575 | 575 | 575 | 575 | 575 | 575 |
| R-squared | 0.29 | 0.82 | 0.37 | 0.82 | 0.41 | 0.83 | -0.82 | 0.83 | 0.41 | 0.83 |
| Number of Countries | 70 | 70 | 60 | 60 | 60 | 60 | 60 | 60 | 60 | 60 |
| Partial R-Squared | | | 0.01 | 0.01 | 0.29 | 0.29 | 0.00 | 0.01 | 0.29 | 0.29 |
| Hansen J p-value | | | 0.46 | 0.42 | 0.69 | 0.56 | 0.51 | 0.28 | 0.42 | 0.44 |
| Kleibergen-Paap F statistic | | | 1.09 | 1.10 | 30.88 | 36.78 | 0.52 | 0.62 | 32.62 | 42.95 |

Notes: Table 7 presents the estimated OLS and second-stage instrumental variables regression coefficients for female employment as the dependent variable for different specifications. POST is instrumented with external debt to GDP ratio, terms of trade, and 5-year change in democracy score in columns (3) and (4). In columns (5) and (6) the instruments also include lagged POST. POST is instrumented with external debt to GDP ratio interacted with the change in world interest rates (proxied by the US Treasury bill rate), terms of trade, and 5-year change in democracy score in columns (7) and (8). In columns (9) and (10) the instruments also include lagged POST. Standard errors clustered at the country level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

## 6. Concluding Remarks

We have argued in this study that there is need for serious research on the impact of structural reforms on a comprehensive list of labor-market outcomes. This is what we have tried to do in this study. We have also argued that micro-level studies might not always be of great value to policy makers when they have to decide on broad policy reforms. Therefore, in this study we look at the impact of structural reforms on macro-level labor-market outcomes, namely the unemployment rate, the employment level, average wage index, labor force participation rates (overall and female) and female employment at the country level.

To our knowledge, there are only two major cross-country empirical studies that look at the impact of trade policy on unemployment rates. One is the paper by Dutt, Mitra and Ranjan (2009) and the other is by Felbermayr, Prat and Schmerer (2011a). Both papers show that countries that have less protectionist (more open) trade policies have lower unemployment rates. This is true both without any controls and after controlling for other policies that have a more direct impact on labor markets. Dutt, Mitra and Ranjan (2009) also find that the short-run impact of structural reforms is an increase in the unemployment rate followed by a reduction in the long run to a lower steady-state unemployment rate.

As mentioned in the introduction, there are several differences between our work and earlier work in the existing literature. Unlike earlier work, our study looks at a multitude of labor-market outcomes. Secondly, unlike the previous studies, we include sample periods in our panel, such that for each country we cover up to ten years before the reforms and ten years after, which gives us greater confidence in our results. Thirdly, our focus is not limited to trade policies and we look at the impact of structural reforms in general (of which trade reform is just one component) on labor-market outcomes. Finally, we are able to use credible instrumental variables for our structural reform variable. These instruments are based on the theoretical literature on the political economy of structural reforms and have been rigorously tested econometrically for their quality.

We have documented the trends on average (across countries) in our labor-market variables around the reform year for each country. This is done by controlling for country fixed effects and in another specification additionally controlling for real GDP, the labor force participation rates and the working age population. We have also run fixed-effects ordinary least squares as well as instrumental variables regressions of our labor-market outcomes on a reform dummy variable, a time trend, square of the time trend and the set of controls mentioned above.

Overall, we find that structural reforms lead to positive outcomes for labor, particularly for informal workers, which is in sharp contrast to the widely held belief that reforms destroy jobs, increase inequality, make the rich richer and do not do much for the poor. Because we have controlled for the time trend and real GDP, our results show the impact of structural reforms on labor-market outcomes beyond what happens through its impact on growth. Redistributive effects in favor of workers, along the lines of the Stolper-Samuelson effect, may be at work.

## References


Alesina, Alberto & Drazen, Allan (1991) "Why are Stabilizations Delayed?," *American Economic Review* 81(5), 1170-88.

Baldwin, Robert (2003) "Openness and Growth: What's the Empirical Relationship?" NBER Working Paper No. 9578, The National Bureau of Economic Research, Cambridge, MA.

Bassanini, Andrea, and Romain Duval (2006) "Employment Patterns in OECD Countries: Reassessing the Roles of Policies and Institutions," OECD Economics Department Working Paper No. 5864.

Bassanini, Andrea, and Romain Duval (2009) "Unemployment, institutions and reform complementarities: Reassessing the Aggregate Evidence for OECD countries," *Oxford Review of Economic Policy* 25, 40-59.

Blanchard, Olivier, and Justin Wolfers (2000) "The Role of Shocks and Institutions in the Rise of European Unemployment: The Aggregate Evidence," *Economic Journal* 110 (462), 1-33.

Boulhol, Herve (2008) "Unemployment and Interactions Between Trade and Labour Market Institutions," CES Working Papers.

Coates, Daniel E. and Ludema, Rodney D. (2001) "A Theory of Trade Policy Leadership," *Journal of Development Economics*, 65(1): 1-29.

Conference Board (2012). Total Economy Database, New York, NY.

Davidson, Carl; Martin, Lawrence and Matusz, Steven (1999) "Trade and search-generated unemployment," *Journal of International Economics* 48(2), 271-299.

Drazen, Allan & Grilli, Vittorio (1993) "The Benefit of Crises for Economic Reforms," American Economic Review 83(3): 598-607.

Dutt, Pushan; Mitra, Devashish and Ranjan, Priya (2009) "International Trade and Unemployment: Theory and Cross-National Evidence," *Journal of International Economics* 78(1), June 2009, 32-44.

Felbermayr, Gabriel; Prat, Julien and Schmerer, Hans-Joerg (2011) "Globalization and labor market outcomes: Wage bargaining, search frictions, and firm heterogeneity," *Journal of Economic Theory* 146 (1), 39-73.

Felbermayr, Gabriel; Prat, Julien and Schmerer, Hans-Joerg (2011a) "Trade and Unemployment: What Do the Data Say?," *European Economic Review* 55 (6), 741-758.

Fernandez, Raquel and Rodrik, Dani (1991). "Resistance to Reform: Status Quo Bias in the Presence of Individual-Specific Uncertainty," American Economic Review 81(5): 1146-55.

Gurr, Ted Robert, Monty G. Marshall, and Keith Jaggers (2012). Political Regime Characteristics and Transitions, 1800-2010.

Hasan, Rana; Mitra, Devashish; Ranjan, Priya and Ahsan, Reshad (2012) "Trade Liberalization and Unemployment: Evidence from India," *Journal of Development Economics* 97(2), 269-280.

Helpman, Elhanan and Itskhoki, Oleg (2010) "Labor market rigidities, trade and unemployment," *Review of Economic Studies* 77, 1100-1137.

Heston, Alan, Robert Summers and Bettina Aten (2011). Penn World Table Version 7.0, Center for International Comparisons of Production, Income and Prices at the University of Pennsylvania, May 2011.

International Monetary Fund (2012). International Financial Statistics, Washington, D.C.

Krishna, Pravin and Mitra, Devashish (2005) “Reciprocated Unilateralism in Trade Policy,” *Journal of International Economics* 65(2), March 2005, 461-487.

Krishna, Pravin and Mitra, Devashish (2008) “Reciprocated Unilateralism in Trade Reforms with Majority Voting,” *Journal of Development Economics* 85(1), 81-93.

Krugman, Paul R. (1993) “What do Undergrads Need to Know About Trade?” *American Economic Review* 83(2), 23-26.

Milner, H. V., and K. Kubota (2005). “Why the Move to Free Trade? Democracy and trade policy in developing countries,” *International Organization* 59(1), 107-43.

Mitra, Devashish; Thomakos, Dimitrios D. and Ulubasoglu, Mehmet A. (2002). “Protection For Sale in a Developing Country: Democracy vs. Dictatorship,” *Review of Economics and Statistics* 84(3): 497-508.

Moore, Marc and Ranjan, Priya (2005) “Globalization vs. Skill-Biased Technological Change, “*Economic Journal* 115(503), 391-422.

Nickell, Stephen, Luca Nunziata, and Wolfgang Ochel (2005) “Unemployment in the OECD Since the 1960s: What do We Know?,” *Economic Journal* 115 (500), 1-27.

Organization for Economic Co-operation and Development (2012). Labor Force Statistics, Paris, France.

Rodrik, Dani (1999) *Has Globalization Gone Too Far?* Peterson Institute of International Economics, Washington DC.

Sachs, Jeffrey D. and Andrew Warner (1995). “Economic Reform and the Process of Global Integration.” *Brookings Papers on Economic Activity* 1: 1–118.

Scarpetta, Stefano (1996) “Assessing the Role of Labour Market Policies and Institutional Settings on Unemployment: A Cross-Country Study,” *OECD Economic Studies* 26, 43-98.

Stock, J. H., and M. Yogo (2005). “Testing for Weak Instruments in Linear IV Regression,” in J. H. Sotkc and D. W. K. Andrews (Eds.), Identification and Inference for Econometric Models: Essays in Honors of Thomas J Rothenberg, New York: Cambridge University Press.

Trefler, Daniel (2004) “The Long and Short of the Canada-U.S. Free Trade Agreement,” *American Economics Review* 94, 870-895.

Wacziarg, Romain and Karen Horn Welch (2008). “Trade Liberalization and Growth: New Evidence.” *The World Bank Economic Review* 22(2): 187-231.

World Bank (2012). World Development Indicators, Washington, D.C.

# Appendix

## a. Data

The macro-level regression analysis uses data on the dates of structural reforms and the country sample from Wacziarg and Welch (2008), which proxies for the year when countries reached a threshold of broad economic reform, such as macroeconomic stabilization, privation, trade opening, and the end of interventionist states, such as communism. The sample consists of 88 countries, the majority being developing countries since most developed countries had already reformed during the sample period.

Six labor market outcomes are considered separately as dependent variables of interest. The unemployment rate is defined as the percentage of the labor force that is without work but available for and seeking employment. The series was constructed using data from the International Monetary Fund *International Financial Statistics* (IFS), the World Bank *World Development Indicators* (WDI), and the International Labour Organization (ILO) *Key Indicators of the Labour Market*, the OECD *Labor Force Statistics*, and other regional agence and country-specific sources. Employment measured in millions is accessed from The Conference Board *Total Economy Database* (TED), and includes employees, the self-employed, unpaid family members who are economically engaged, apprentices, and the military. Employment series for countries not available from the TED were accessed from the WDI as total workers aged 15 and older. Female employment measured in millions is accessed from the WDI and represents women workers aged 15 and older. The wage index, accessed from the IFS, is an index of wage earnings with 2005 equal to 100 as the base year.

The labor force participation rate, accessed from the WDI, is the labor force as a percentage of the working-age population. Total labor force comprises people ages 15 and older who meet the ILO definition of the economically active population: all people who supply labor for the production of goods and services during a specified period. It includes both employed and unemployed. While national practices vary in the treatment of such groups as the armed forces and seasonal or part-time workers, in general the labor force includes the armed forces, the unemployed, and first-time job-seekers but excludes homemakers and other unpaid caregivers and informal workers. Labor force participation rate series for countries not available from the WDI were accessed from the IFS as the labor force as a percentage of the population aged 15 and older. The female labor force participation rate, accessed from the WDI, is women workers as a percentage of the female working age population.

## b. Countries in dataset and year of liberalization

**Table A.1: Countries in Dataset and Year of Liberalization**

| Country | Reform Year | Unemployment Rate | Employment | Wage Index | LFP Rate | Female LFP Rate | Female Employment |
|---|---|---|---|---|---|---|---|
| Albania | 1992 | X | X | X | X | X | X |
| Algeria | Always Closed | | | | | | |
| Angola | Always Closed | | | | | | |
| Argentina | 1991 | X | X | | X | X | X |
| Armenia | 1995 | X | X | X | X | X | X |
| Australia | 1964 | X | X | X | | | |
| Austria | 1960 | | X | X | X | | |
| Azerbaijan | 1995 | X | X | | X | X | X |
| Bangladesh | 1996 | X | X | | X | X | X |
| Barbados | 1966 | | X | | | | |
| Belarus | Always Closed | | | | | | |
| Belgium | Always Open | | | | | | |
| Benin | 1990 | X | X | | X | X | X |
| Bolivia | 1985 | X | X | | X | X | X |
| Botswana | 1979 | | | | X | X | |
| Brazil | 1991 | X | X | | X | X | X |
| Bulgaria | 1991 | X | X | | X | X | X |
| Burkina Faso | 1998 | X | X | | X | X | X |
| Burundi | 1999 | | X | | X | X | X |
| Cameroon | 1993 | X | X | | X | X | X |
| Canada | Always Open | | | | | | |
| Cape Verde | 1991 | | X | | X | X | X |
| Central African Republic | Always Closed | | | | | | |
| Chad | Always Closed | | | | | | |
| Chile | 1976 | X | X | | X | X | |
| China | Always Closed | | | | | | |
| Colombia | 1986 | X | X | | X | X | X |
| Congo, Dem. Rep. | Always Closed | | | | | | |
| Congo, Rep. | Always Closed | | | | | | |
| Costa Rica | 1986 | X | X | | X | X | X |
| Cote d'Ivoire | 1994 | X | X | | X | X | X |
| Croatia | Always Closed | | | | | | |
| Cyprus | 1960 | | X | | | | |
| Czech Republic | 1991 | X | X | X | X | X | X |
| Denmark | Always Open | | | | | | |
| Dominican Republic | 1992 | X | X | | X | X | X |
| Ecuador | 1991 | X | X | | X | X | X |
| Egypt, Arab Rep. | 1995 | X | X | | X | X | X |
| El Salvador | 1989 | X | X | | X | X | X |
| Estonia | Always Closed | | | | | | |
| Ethiopia | 1996 | X | X | | X | X | X |
| Finland | 1960 | X | X | | | | |
| France | Always Open | | | | | | |
| Gabon | Always Closed | | | | | | |
| Gambia, The | 1985 | | X | | X | X | X |
| Georgia | 1996 | X | X | | X | X | X |
| Germany | Always Open | | | | | | |
| Ghana | 1985 | | X | | X | X | X |
| Greece | Always Open | | | | | | |
| Guatemala | 1988 | X | X | | X | X | X |
| Guinea | 1986 | | X | | X | X | X |
| Guinea-Bissau | 1987 | | X | | X | X | X |
| Guyana | 1988 | X | X | | X | X | X |
| Haiti | Always | | | | | | |

| | | | | | | | |
|---|---|---|---|---|---|---|---|
| | Closed | | | | | | |
| Honduras | 1991 | X | X | | X | X | X |
| | Always | | | | | | |
| Hong Kong, China | Open | | | | | | |
| Hungary | 1990 | X | X | X | X | X | X |
| | Always | | | | | | |
| India | Closed | | | | | | |
| Indonesia | 1970 | X | X | | | | |
| | Always | | | | | | |
| Iran | Closed | | | | | | |
| | Always | | | | | | |
| Iraq | Closed | | | | | | |
| Ireland | 1966 | | X | | | | |
| Israel | 1985 | X | X | X | X | X | X |
| | Always | | | | | | |
| Italy | Open | | | | | | |
| Jamaica | 1989 | X | X | | X | X | X |
| Japan | 1964 | X | X | X | | | |
| Jordan | 1965 | | | | | | |
| | Always | | | | | | |
| Kazakhstan | Closed | | | | | | |
| Kenya | 1993 | | X | | X | X | X |
| Korea, Rep. | 1968 | X | X | | | | |
| Kyrgyz Republic | 1994 | X | X | X | X | X | X |
| Latvia | 1993 | X | X | X | X | X | X |
| | Always | | | | | | |
| Lesotho | Closed | | | | | | |
| | Always | | | | | | |
| Liberia | Closed | | | | | | |
| Lithuania | 1993 | X | X | X | X | X | X |
| | Always | | | | | | |
| Luxembourg | Open | | | | | | |
| Macedonia, FYR | 1994 | X | | X | X | X | X |
| Madagascar | 1996 | X | X | | X | X | X |
| | Always | | | | | | |
| Malawi | Closed | | | | | | |
| Malaysia | 1963 | | X | | | | |
| Mali | 1988 | | X | | X | X | X |
| | Always | | | | | | |
| Malta | Closed | | | | | | |
| Mauritania | 1995 | X | X | | X | X | X |
| Mauritius | 1968 | | | | | | |
| Mexico | 1986 | X | X | X | X | X | X |
| Moldova | 1994 | X | X | | X | X | X |
| Morocco | 1984 | X | X | | X | X | X |
| Mozambique | 1995 | | X | | X | X | X |
| | Always | | | | | | |
| Myanmar | Closed | | | | | | |
| Nepal | 1991 | X | X | | X | X | X |
| | Always | | | | | | |
| Netherlands | Open | | | | | | |
| New Zealand | 1986 | X | X | X | X | X | X |
| Nicaragua | 1991 | X | X | | X | X | X |
| Niger | 1994 | X | X | | X | X | X |
| | Always | | | | | | |
| Nigeria | Closed | | | | | | |
| | Always | | | | | | |
| Norway | Open | | | | | | |
| Pakistan | 2001 | X | X | | X | X | X |
| Panama | 1996 | X | X | | X | X | X |
| | Always | | | | | | |
| Papua New Guinea | Closed | | | | | | |
| Paraguay | 1989 | X | X | | X | X | X |
| Peru | 1991 | X | X | | X | X | X |
| Philippines | 1988 | X | X | | X | X | X |
| Poland | 1990 | X | X | X | X | X | X |
| | Always | | | | | | |
| Portugal | Open | | | | | | |
| Romania | 1992 | X | X | X | X | X | X |
| | Always | | | | | | |
| Russian Federation | Closed | | | | | | |
| | Always | | | | | | |
| Rwanda | Closed | | | | | | |
| | Always | | | | | | |
| Senegal | Closed | | | | | | |
| Sierra Leone | 2001 | | X | | X | X | X |
| Singapore | 1965 | X | X | | | | |
| Slovak Republic | 1991 | X | X | X | X | X | X |
| Slovenia | 1991 | X | X | X | X | X | X |
| Somalia | Always | | | | | | |

| | | | | | | | |
|---|---|---|---|---|---|---|---|
| | Closed | | | | | | |
| South Africa | 1991 | X | X | | X | X | X |
| Spain | Always Open | | | | | | |
| Sri Lanka | 1991 | X | X | X | X | X | X |
| Swaziland | Always Closed | | | | | | |
| Sweden | 1960 | X | X | X | | | |
| Switzerland | Always Open | | | | | | |
| Syrian Arab Republic | Always Closed | | | | | | |
| Taiwan, China | 1963 | | X | | | | |
| Tajikistan | 1996 | | X | | X | X | X |
| Tanzania | 1995 | X | X | | X | X | X |
| Thailand | Always Open | | | | | | |
| Togo | Always Closed | | | | | | |
| Trinidad and Tobago | 1992 | X | X | X | X | X | X |
| Tunisia | 1989 | X | X | | X | X | X |
| Turkey | 1989 | X | X | | X | X | X |
| Turkmenistan | Always Closed | | | | | | |
| Uganda | 1988 | | X | | X | X | X |
| Ukraine | Always Closed | | | | | | |
| United Kingdom | Always Open | | | | | | |
| United States | Always Open | | | | | | |
| Uruguay | 1990 | X | X | | X | X | X |
| Uzbekistan | Always Closed | | | | | | |
| Venezuela | 1996 | X | X | | X | X | X |
| Yemen, Rep. | Always Open | | | | | | |
| Yugoslavia, FR (Serbia/Montenegro) | 2001 | X | X | | | | |
| Zambia | 1993 | X | X | | X | X | X |
| Zimbabwe | Always Closed | | | | | | |

Notes: Tale A.1 presents the sample of countries from Wacziard and Welch (2008), each country's liberalization date in the first column, and an indicator if the country is included in the sample for each dependent variable of interest in the second through seventh columns.

## c. First-stage regression results

**Table A.2: Unemployment Rate**

| VARIABLES | (3)<br>IV-FE | (4)<br>IV-FE | (5)<br>IV-FE | (6)<br>IV-FE | (7)<br>IV-FE | (8)<br>IV-FE | (9)<br>IV-FE | (10)<br>IV-FE |
|---|---|---|---|---|---|---|---|---|
| Time | 0.07*** | 0.07*** | 0.03*** | 0.03*** | 0.07*** | 0.07*** | 0.03*** | 0.03*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| Timesq | -0.00** | -0.00** | -0.00*** | -0.00*** | -0.00 | -0.00* | -0.00*** | -0.00*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| WorkingAgePop | | -0.00 | | 0.00 | | -0.00 | | 0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| rGDP_PWT | | 0.00 | | -0.00 | | 0.00 | | -0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| LFPRate | | 0.00 | | 0.00 | | 0.00 | | 0.00 |
| | | (0.01) | | (0.01) | | (0.01) | | (0.01) |
| IV1B_Debt | 0.05*** | 0.05*** | 0.02** | 0.02** | | | | |
| | (0.02) | (0.02) | (0.01) | (0.01) | | | | |
| IV2_TofT | 0.00* | 0.00** | 0.00** | 0.00* | 0.00* | 0.00* | 0.00* | 0.00* |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| IV4A_Dem5 | -0.00 | -0.00 | -0.00 | -0.00 | -0.00 | -0.00 | -0.00 | -0.00 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| LPOST | | | 0.63*** | 0.63*** | | | 0.64*** | 0.64*** |
| | | | (0.02) | (0.02) | | | (0.02) | (0.02) |
| IV1A_Debt | | | | | 0.01 | 0.01 | -0.00 | -0.00 |
| | | | | | (0.01) | (0.01) | (0.01) | (0.01) |
| Observations | 514 | 514 | 514 | 514 | 514 | 514 | 514 | 514 |
| R-squared | 0.71 | 0.71 | 0.83 | 0.83 | 0.71 | 0.71 | 0.83 | 0.83 |
| Number of Countries | 48 | 48 | 48 | 48 | 48 | 48 | 48 | 48 |

Notes: Table A.2 presents the estimated first-stage instrumental variables regression coefficients for unemployment rate as the dependent variable for different specifications. POST is instrumented with external debt to GDP ratio, terms of trade, and 5-year change in democracy score in columns (3) and (4). In columns (5) and (6) the instruments also include lagged POST. POST is instrumented with external debt to GDP ratio interacted with the change in world interest rates (proxied by the US Treasury bill rate), terms of trade, and 5-year change in democracy score in columns (7) and (8). In columns (9) and (10) the instruments also include lagged POST. Standard errors clustered at the country level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

## Table A.3: Employment

| VARIABLES | (3) IV-FE | (4) IV-FE | (5) IV-FE | (6) IV-FE | (7) IV-FE | (8) IV-FE | (9) IV-FE | (10) IV-FE |
|---|---|---|---|---|---|---|---|---|
| Time | 0.07*** | 0.07*** | 0.03*** | 0.03*** | 0.07*** | 0.07*** | 0.03*** | 0.03*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| Timesq | -0.00*** | -0.00*** | -0.00*** | -0.00*** | -0.00*** | -0.00*** | -0.00*** | -0.00*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| WorkingAgePop | | -0.00 | | 0.00 | | -0.00 | | 0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| rGDP_PWT | | 0.00 | | -0.00 | | 0.00 | | -0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| LFPRate | | -0.01 | | -0.00 | | -0.01 | | -0.00 |
| | | (0.01) | | (0.00) | | (0.01) | | (0.00) |
| IV1B_Debt | 0.05 | 0.05 | -0.00 | -0.00 | | | | |
| | (0.05) | (0.05) | (0.02) | (0.02) | | | | |
| IV2_TofT | 0.00 | 0.00 | 0.00* | 0.00 | 0.00 | 0.00 | 0.00* | 0.00 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| IV4A_Dem5 | -0.00 | -0.01 | -0.00 | -0.00 | -0.01 | -0.01 | -0.00 | -0.00 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| LPOST | | | 0.61*** | 0.61*** | | | 0.61*** | 0.61*** |
| | | | (0.01) | (0.01) | | | (0.02) | (0.02) |
| IV1A_Debt | | | | | 0.01 | 0.01 | 0.01 | 0.01 |
| | | | | | (0.01) | (0.01) | (0.01) | (0.01) |
| Observations | 724 | 723 | 724 | 723 | 724 | 723 | 724 | 723 |
| R-squared | 0.68 | 0.68 | 0.81 | 0.81 | 0.68 | 0.68 | 0.81 | 0.81 |
| Number of Countries | 60 | 60 | 60 | 60 | 60 | 60 | 60 | 60 |

Notes: Table A.3 presents the estimated first-stage instrumental variables regression coefficients for employment as the dependent variable for different specifications. POST is instrumented with external debt to GDP ratio, terms of trade, and 5-year change in democracy score in columns (3) and (4). In columns (5) and (6) the instruments also include lagged POST. POST is instrumented with external debt to GDP ratio interacted with the change in world interest rates (proxied by the US Treasury bill rate), terms of trade, and 5-year change in democracy score in columns (7) and (8). In columns (9) and (10) the instruments also include lagged POST. Standard errors clustered at the country level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

## Table A.4: Wage Index

| VARIABLES | (3)<br>IV-FE | (4)<br>IV-FE | (5)<br>IV-FE | (6)<br>IV-FE | (7)<br>IV-FE | (8)<br>IV-FE | (9)<br>IV-FE | (10)<br>IV-FE |
|---|---|---|---|---|---|---|---|---|
| Time | 0.07*** | 0.06*** | 0.03*** | 0.03*** | 0.07*** | 0.06*** | 0.03*** | 0.03*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.01) | (0.01) | (0.01) |
| Timesq | -0.00*** | -0.00** | -0.00*** | -0.00*** | -0.00 | -0.00 | -0.00** | -0.00*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| WorkingAgePop | | 0.00** | | 0.00** | | 0.00 | | 0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| rGDP_PWT | | -0.00** | | -0.00** | | -0.00 | | -0.00** |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| LFPRate | | -0.01 | | 0.00 | | -0.01 | | 0.00 |
| | | (0.02) | | (0.01) | | (0.01) | | (0.01) |
| IV1B_Debt | 0.81** | 1.11** | 0.00 | 0.22 | | | | |
| | (0.28) | (0.41) | (0.21) | (0.21) | | | | |
| IV2_TofT | 0.00*** | 0.00 | 0.00*** | 0.00 | 0.00** | 0.00 | 0.00 | 0.00 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| IV4A_Dem5 | 0.01 | 0.01 | -0.00 | -0.00 | 0.01 | 0.01 | -0.00 | -0.00 |
| | (0.01) | (0.01) | (0.00) | (0.00) | (0.01) | (0.01) | (0.01) | (0.01) |
| LPOST | | | 0.64*** | 0.60*** | | | 0.60*** | 0.61*** |
| | | | (0.01) | (0.04) | | | (0.11) | (0.09) |
| IV1A_Debt | | | | | 0.11 | 0.10 | 0.06 | 0.06 |
| | | | | | (0.07) | (0.08) | (0.07) | (0.07) |
| Observations | 66 | 66 | 66 | 66 | 66 | 66 | 66 | 66 |
| R-squared | 0.75 | 0.77 | 0.84 | 0.84 | 0.74 | 0.75 | 0.84 | 0.85 |
| Number of Countries | 9 | 9 | 9 | 9 | 9 | 9 | 9 | 9 |

Notes: Table A.4 presents the estimated first-stage instrumental variables regression coefficients for wage index as the dependent variable for different specifications. POST is instrumented with external debt to GDP ratio, terms of trade, and 5-year change in democracy score in columns (3) and (4). In columns (5) and (6) the instruments also include lagged POST. POST is instrumented with external debt to GDP ratio interacted with the change in world interest rates (proxied by the US Treasury bill rate), terms of trade, and 5-year change in democracy score in columns (7) and (8). In columns (9) and (10) the instruments also include lagged POST. Standard errors clustered at the country level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

## Table A.5: Labor Force Participation Rate

| VARIABLES | (3) IV-FE | (4) IV-FE | (5) IV-FE | (6) IV-FE | (7) IV-FE | (8) IV-FE | (9) IV-FE | (10) IV-FE |
|---|---|---|---|---|---|---|---|---|
| Time | 0.07*** | 0.07*** | 0.03*** | 0.03*** | 0.07*** | 0.07*** | 0.03*** | 0.03*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| Timesq | -0.00 | -0.00 | -0.00*** | -0.00*** | -0.00 | -0.00 | -0.00*** | -0.00*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| WorkingAgePop | | 0.00 | | 0.00 | | 0.00 | | 0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| rGDP_PWT | | 0.00 | | -0.00 | | 0.00 | | -0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| IV1B_Debt | 0.03** | 0.03** | 0.01 | 0.01 | | | | |
| | (0.01) | (0.01) | (0.01) | (0.01) | | | | |
| IV2_TofT | 0.00** | 0.00** | 0.00** | 0.00** | 0.00** | 0.00** | 0.00** | 0.00* |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| IV4A_Dem5 | -0.00 | -0.00 | -0.00 | -0.00 | -0.00 | -0.00 | -0.00 | -0.00 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| LPOST | | | 0.64*** | 0.64*** | | | 0.64*** | 0.64*** |
| | | | (0.01) | (0.01) | | | (0.01) | (0.01) |
| IV1A_Debt | | | | | 0.01 | 0.01 | 0.00 | 0.00 |
| | | | | | (0.01) | (0.01) | (0.01) | (0.01) |
| Observations | 960 | 960 | 960 | 960 | 960 | 960 | 960 | 960 |
| R-squared | 0.73 | 0.73 | 0.84 | 0.84 | 0.73 | 0.73 | 0.84 | 0.84 |
| Number of Countries | 62 | 62 | 62 | 62 | 62 | 62 | 62 | 62 |
| Time | 0.07*** | 0.07*** | 0.03*** | 0.03*** | 0.07*** | 0.07*** | 0.03*** | 0.03*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |

Notes: Table A.5 presents the estimated first-stage instrumental variables regression coefficients for labor force participation rate as the dependent variable for different specifications. POST is instrumented with external debt to GDP ratio, terms of trade, and 5-year change in democracy score in columns (3) and (4). In columns (5) and (6) the instruments also include lagged POST. POST is instrumented with external debt to GDP ratio interacted with the change in world interest rates (proxied by the US Treasury bill rate), terms of trade, and 5-year change in democracy score in columns (7) and (8). In columns (9) and (10) the instruments also include lagged POST. Standard errors clustered at the country level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

## Table A.6: Female Labor Force Participation Rate

| VARIABLES | (3)<br>IV-FE | (4)<br>IV-FE | (5)<br>IV-FE | (6)<br>IV-FE | (7)<br>IV-FE | (8)<br>IV-FE | (9)<br>IV-FE | (10)<br>IV-FE |
|---|---|---|---|---|---|---|---|---|
| Time | 0.07*** | 0.07*** | 0.03*** | 0.03*** | 0.07*** | 0.07*** | 0.03*** | 0.03*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| Timesq | -0.00 | -0.00 | -0.00*** | -0.00*** | -0.00 | -0.00 | -0.00*** | -0.00*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| WorkingAgePop | | 0.00 | | 0.00 | | 0.00 | | 0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| rGDP_PWT | | 0.00 | | -0.00 | | 0.00 | | -0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| IV1B_Debt | 0.03** | 0.03** | 0.01 | 0.01 | | | | |
| | (0.01) | (0.01) | (0.01) | (0.01) | | | | |
| IV2_TofT | 0.00** | 0.00** | 0.00** | 0.00** | 0.00** | 0.00** | 0.00** | 0.00* |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| IV4A_Dem5 | -0.00 | -0.00 | -0.00 | -0.00 | -0.00 | -0.00 | -0.00 | -0.00 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| LPOST | | | 0.64*** | 0.64*** | | | 0.64*** | 0.64*** |
| | | | (0.01) | (0.01) | | | (0.01) | (0.01) |
| IV1A_Debt | | | | | 0.01 | 0.01 | 0.00 | 0.00 |
| | | | | | (0.01) | (0.01) | (0.01) | (0.01) |
| Observations | 960 | 960 | 960 | 960 | 960 | 960 | 960 | 960 |
| R-squared | 0.73 | 0.73 | 0.84 | 0.84 | 0.73 | 0.73 | 0.84 | 0.84 |
| Number of Countries | 62 | 62 | 62 | 62 | 62 | 62 | 62 | 62 |
| Time | 0.07*** | 0.07*** | 0.03*** | 0.03*** | 0.07*** | 0.07*** | 0.03*** | 0.03*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |

Notes: Table A.6 presents the estimated first-stage instrumental variables regression coefficients for female labor force participation rate as the dependent variable for different specifications. POST is instrumented with external debt to GDP ratio, terms of trade, and 5-year change in democracy score in columns (3) and (4). In columns (5) and (6) the instruments also include lagged POST. POST is instrumented with external debt to GDP ratio interacted with the change in world interest rates (proxied by the US Treasury bill rate), terms of trade, and 5-year change in democracy score in columns (7) and (8). In columns (9) and (10) the instruments also include lagged POST. Standard errors clustered at the country level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

## Table A.7: Female Employment

| VARIABLES | (3) IV-FE | (4) IV-FE | (5) IV-FE | (6) IV-FE | (7) IV-FE | (8) IV-FE | (9) IV-FE | (10) IV-FE |
|---|---|---|---|---|---|---|---|---|
| Time | 0.10*** | 0.11*** | 0.05*** | 0.06*** | 0.11*** | 0.11*** | 0.05*** | 0.06*** |
| | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) |
| Timesq | -0.01*** | -0.01*** | -0.00*** | -0.00*** | -0.01*** | -0.01*** | -0.00*** | -0.00*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| WorkingAgePop | | -0.00 | | -0.00 | | -0.00 | | -0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| rGDP_PWT | | 0.00 | | -0.00 | | 0.00 | | -0.00 |
| | | (0.00) | | (0.00) | | (0.00) | | (0.00) |
| LFPRate | | -0.00 | | -0.00 | | -0.00 | | -0.00 |
| | | (0.01) | | (0.01) | | (0.01) | | (0.01) |
| IV1B_Debt | 0.06 | 0.05 | 0.02 | 0.01 | | | | |
| | (0.04) | (0.04) | (0.02) | (0.03) | | | | |
| IV2_TofT | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| IV4A_Dem5 | 0.00 | 0.00 | -0.00 | 0.00 | -0.00 | -0.00 | -0.00 | 0.00 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| LPOST | | | 0.48*** | 0.48*** | | | 0.49*** | 0.48*** |
| | | | (0.05) | (0.04) | | | (0.05) | (0.04) |
| IV1A_Debt | | | | | 0.00 | 0.00 | 0.00 | 0.00 |
| | | | | | (0.01) | (0.01) | (0.01) | (0.01) |
| Observations | 575 | 575 | 575 | 575 | 575 | 575 | 575 | 575 |
| R-squared | 0.65 | 0.65 | 0.75 | 0.75 | 0.64 | 0.65 | 0.75 | 0.75 |
| Number of Countries | 60 | 60 | 60 | 60 | 60 | 60 | 60 | 60 |

Notes: Table A.7 presents the estimated first-stage instrumental variables regression coefficients for female employment as the dependent variable for different specifications. POST is instrumented with external debt to GDP ratio, terms of trade, and 5-year change in democracy score in columns (3) and (4). In columns (5) and (6) the instruments also include lagged POST. POST is instrumented with external debt to GDP ratio interacted with the change in world interest rates (proxied by the US Treasury bill rate), terms of trade, and 5-year change in democracy score in columns (7) and (8). In columns (9) and (10) the instruments also include lagged POST. Standard errors clustered at the country level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

# Chapter 3

# Measuring Capital Matters

**Claire H. Hollweg**
**with Daniel Lederman**

# 1. Introduction and Motivation

Estimates of Brazil's productivity trends are puzzling. Macroeconomic estimates of Brazil's Total Factor Productivity (TFP), or the portion of growth of GDP that is not due to labor or capital accumulation, document a negative growth rate of TFP throughout the 1980s and into the early 1990s, with a slight increase in TFP after 1992.[20] If we take these estimates of Brazilian productivity seriously, we would conclude not only that the trade and economic reforms (which included price stabilization) implemented in the late 1980s did not have the expected effects on productivity growth, but that the economy of Brazil is less efficient in the 21st century than in the early 1980s. However, microeconomic estimates of firm-level productivity of Brazil document increases in both within-firm and within-industry productivities after the reforms.[21] This study offers a potential explanation for Brazil's productivity puzzle by focusing on one issue related to the measurement of macroeconomic estimates of TFP: the role of the price deflator.

This study argues that Brazil's macroeconomic estimates of TFP are mis-measured due to mis-measurement of the price (and thus stock) of capital. We correct for this mis-measurement by constructing new capital price indexes (following three definitions of capital goods used in the literature). We proxy for the price of capital using unit values from international trade data on capital goods and adjust the index to reflect domestic Brazilian prices using the real (R$)/US dollar ($) nominal exchange rate and Brazilian tariff data. With the capital price index in hand, we are able to appropriately deflate the gross fixed capital formation series published in the Brazilian national accounts, which is subsequently used to compute a new series of the capital stock employing the perpetual inventory method (PIM). Our newly constructed capital stock series are then used to replicate simple estimates of TFP in a standard growth accounting framework.

[20] See, for example, Bugarin et al. (2002), Gomes, Pessôa, and Veloso (2003), Silva Filho (2001), and Pinheiro et al. (2001).
[21] See, for example, Muendler, Servén and Sepúlveda (2004).

The GDP deflator is commonly used to deflate the investment series prior to constructing a series of unobservable capital stocks using the PIM. This implicitly assumes that the price of fixed capital investments is equal to the price of all goods and services in the economy. However, mis-measurement caused by the divergence in these relative prices is often ignored. If the price of capital investments rises more than the price of all goods and services in the economy, the divergence in these relative prices would result in an overestimated capital stock and underestimated TFP. As pointed out by Pritchett (2000), this is particularly problematic when the data cover periods of economic reforms, especially trade reforms and price-stabilization regimes. During periods of high inflation and frequent and large nominal exchange rate variations, the relative prices of all tradable goods can vary substantially with respect to the prices of non-tradable goods and services. Since the latter are part of GDP, whereas capital goods can be safely assumed to be tradable (even if produced domestically), to estimate the stock of capital through the accumulation of investments deflated by the price of GDP would overestimate the value of this stock and underestimate TFP. It is very likely that the dramatic drop in the estimates of Brazil's TFP during the late 1980s and early 1990s was due to the "capital stock inflation effect," or the mis-measurement of the relative price of capital over the GDP deflator.

However, the relative price of capital may have dropped after the trade reforms as a result of lower tariffs. This "trade policy effect" would imply that the accumulation of capital between 1988 and 1995 (when the trade reforms under MERCOSUR was re-invigorated) might be underestimated and thus TFP could be overestimated. But Brazilian trade policies are notorious for protecting capital goods producers, even after 1994 under the MERCOSUR Customs Union.

Our results document a significant divergence between the newly constructed capital price indexes and Brazil's published price deflators. The percentage difference between the capital price indexes and the GDP deflator ranges from 23 to 128 percent, and for the fixed capital formation price deflator from 48 to 173 percent. When decomposing the capital price indexes, we see that the "trade-policy effect" is not large enough to compensate for the "capital stock inflation effect." Our results also document a fall in Brazil's capital-output ratio, as opposed to significant capital deepening observed in the Brazilian national accounts during this time period. Finally, our results show a significant recovery in Brazil's TFP between 1992 and 2006, with a cumulative increase ranging between 22 and 30 percent. This amounts to an average annual growth rate in TFP ranging between 1.5 and 1.9 percent. Overall, the level of TFP in 2006 is between 8 and 15 percent higher than in 1989. However, due to the limited years for which trade data exist, we acknowledge that we are not be able to explain the downfall of TFP in the 1980s, only its recovery.

The remainder of the chapter proceeds as follows. Section 2 provides a brief literature review of studies that document and attempt to shed light on Brazil's TFP trends. Section 3 details how capital goods are identified in trade data. Section 4 provides a detailed overview of the methodology used to construct the capital price index and Section 5 the index results. Section 6 presents the methodology to measure the new series of the capital stock and TFP as well as the results for Brazil for 1989 to 2006. Section 7 concludes.

## 2. Related Literature

Multiple studies document a fall in Brazil's aggregate TFP measures after the trade and economic reforms implemented in Brazil in the late 1980s.[22] Each of these studies calculates TFP from estimates of the capital stock, hours worked, and an aggregate production function. However, differences arise within the literature in terms of data sources for the capital series. Some studies use investment data from Penn World Tables to construct a capital stock series using the PIM. Commonly used is data reported by the Instituto de Pesquisa Econômica

[22] This trend is not unique to Brazil. Ferreira, Pessôa, and Veloso (2006) show that TFP levels in Latin America as a whole declined since the mid-seventies until 2000, despite high and increasing productivity levels relative to the United States and other regions prior to this period.

Aplicada (IPEA). The IPEA data provides a series of gross fixed capital for Brazil from 1908 to 1970 in 1980 prices. Morandi and Reis (2003) updated and extended the series until 2000, and also constructed a capital stock series using the PIM (see Morandi (1998) and Morandi and Reis (2003)). This study revisits work similar to Bugarin et al. (2002), Gomes, Pessôa and Veloso (2003), and Eller-Jr. (2012).

Bugarin et al. (2002) assume Hicks-neutral technological change and a Cobb-Douglas production function to decompose changes in Brazil's GDP due to growth of TFP, changes in the capital intensity, and changes in hours worked per working-age person from 1980-1998 using a standard growth accounting framework. From 1980-1988, representing a stagnation in Brazil, despite a 1.06 percent increase in output per working age person, TFP accounted for -2.41 percent of this change. From 1988-1998, representing a depression in Brazil, output per working age person feel by 0.34 percent, while TFP accounted for -0.85 percent of this change. Furthermore, the authors use the neoclassical growth model to show that the behavior of the Brazilian economy in the 1980s and 1990s may be explained by exogenous productivity shocks, and are able to explain two thirds of the decline in TFP by such shocks.

Gomes, Pessôa, and Veloso (2003) analyze the evolution of TFP for the Brazilian economy from 1950 to 2000 using a Cobb-Douglas production function with Hicks-neutral and Harrod-neutral technological change. The Harrod-neutral, or labor augmenting, technology is assumed to be constant and common to all economies, representing the evolution of the technological frontier. This component is the calibrated growth rate of labor productivity based on long-term behavior of output per workers in the United States. The portion that corresponds to the difference between the evolution of TFP and the technological frontier is called the evolution of discounted total factor productivity, PTFD. The authors interpret this component to be country-specific productivity, versus the technological frontier that corresponds to productivity growth of the economy resulting from links with other market economies. The authors find that Brazil's economy was on a path of balanced growth between

1950 and 1967 with a stable capital-output ratio. Between 1967 and 1976, there were significant increases in TFP, particularly compared to the technological frontier, and a slight fall in the capital-output ratio. Between 1976 and 1992, Brazil's TFP dropped significantly compared to the technological frontier, and Brazil also experienced strong capital deepening. From 1992-2000, there is evidence of a balanced growth path with stability of the capital-output ratio, with the TFP growth rate being determined by the technological frontier.

Eller-Jr. (2012) uses a constant elasticity of substitution production function to separate capital productivity and labor productivity. The author finds that labor productivity is more important in explaining TFP than capital productivity, with the greatest productivity gains from improved labor via human capital. Similar in spirit to this study, Eller-Jr. (2012) also analyzes the role of relative prices in influencing TFP estimates by using a capital price deflator that also takes into account the prices of buildings. Since the prices of buildings increased more than the general price level, not accounting for these price increases overestimates the value of this stock and underestimates TFP. Similarly, Ferreira, Ellery-Jr., and Gomes (2008) show that the capital stock in Brazil is overestimated when constructed using a price index that fails to take into account large increases in the relative price of construction in the late 1980s. The authors deflated capital investments in machinery, equipment, and buildings by a price index particular to each prior to constructing a capital stock series using the PIM. Carvalho Filho and Chamon (2011) also show that other estimates of macroeconomic variables in Brazil are biased stemming from periods of change. Correcting for this bias, the authors find higher GDP growth than official statistics suggest, but the study did not attempt to reconcile biased productivity measures.

## 3. Definition of Capital Goods

The first step in building a capital price index using unit values from international trade data is to identify trade in capital goods. Trade data are available by type of product, but not according to the way in which the product is used, for example as a consumption good versus an investment good. As such, direct measures of trade in investment goods are not available. Instead, to proxy for the price of capital, three classifications of capital goods imports are defined following related literature. Table 1 lists the relative chapters and classes covered for each of the definitions, as well as the literature in which each of these definitions is used. The proceeding analysis is then conducted using each of these three definitions.

**Table 1: Definition of Capital Goods**

| Chapters and Classes Covered | Authors and Papers |
|---|---|
| **United Nations Broad Economic Categories (BEC)** | |
| 41 – Capital goods (except transport equipment)<br>521 – Transport equipment and parts and accessories thereof, industrial | Hiratsuka (2008); and Frensch and Wittich (2009). Also see Turkcan (2007) for defining capital goods along these lines. |
| **Bureau of Economic Analysis (BEA) 34-Industry Code (adapted from the United Nations International Comparison Program)** | |
| 20 – Farm and garden machinery<br>21 – Construction, mining, etc.<br>22 – Computer and office equipment<br>23 – Other nonelectric machinery<br>24 – Household appliances<br>25 – Household audio and video, etc.<br>26 – Electronic components<br>27 – Other electrical machinery<br>33 – Instruments and apparatus | Alfaro and Ahmed (2007); Alfaro and Hammel (2007); and Eaton and Kortum (2001). Adapted from De Long and Summers (1991). |
| **Standard International Trade Classification (SITC)** | |
| 7 – Machinery and transport equipment:<br>71 – Power-generating machinery and equipment<br>72 – Machinery specialized for particular industries<br>73 – Metal machinery<br>74 – General industrial machinery and equipment and machine parts<br>75 – Office machines and automatic data-processing machines<br>76 – Telecommunications and sound-recording and reproducing apparatus and equipment<br>77 – Electrical machinery, apparatus and appliances and electrical parts thereof (including non-electrical counterparts of electrical household-type equipment)<br>78 – Road vehicles (including air-cushion vehicles)<br>79 – Other transport equipment | Bergstrang (1983); Bergstrand (1990); Xu and Wang (1999); and Xu and Wang (2000). |

Sources: Feenstra, Lipsey, and Bowen (1997); United States Department of Commerce (1992); United Nations Statistics Division.
Notes: Table 1 presents the chapters and classes covered for each definition of capital goods, as well as the literature in which each of these definitions is used.

Each of these classifications attempts to capture international trade of investment goods. First, capital goods are classified according to the United Nations Statistics Division's Classification by Broad Economic Categories (BEC). This classification allows for almost all of the basic standard international trade classification (SITC) categories to be grouped into major system of national accounts activities, and includes 471 capital goods categories. However, only a subset of these capital goods categories represents investment versus consumption goods.

Second, capital goods are more simply classified according to Chapter 7 of the SITC. The three-digit SITC classification is typically considered an industry for econometric purposes. Xu and Wang (1999; 2000) use imports of machinery and transport equipment (SITC 7) as a proxy for imports of capital goods, which is also adopted in this study. Bergstrand (I983) uses the two-digit SITC of nonelectrical machinery (71), electrical machinery (72), and transportation equipment (73) to represent manufacturing industries.

Third, capital goods are classified according to the Bureau of Economic Analysis's 34-Industry Code. Eaton and Kortum (2001) approximate trade in capital equipment by trade in goods associated with major equipment producing industries, including output of nonelectrical equipment, electrical equipment, and instruments industries. Output from these industries is more likely to be used for investment rather than consumption. They identify the equipment-producing industries after consulting input output tables and capital flows tables of domestic transactions of three major capital-goods producing countries. These three sectors contribute about two-third of investment goods overall and over three-fourths of investment goods used in manufacturing (Eaton and Kortum 2001). Only about 40 percent of the output of these industries constitutes final investment goods, with the rest used mainly as intermediaries (Eaton and Kortum 2001). However, whether output is final investment or an intermediate is not important in this analysis.

## 4. Construction of the Capital Price Index

After identifying which internationally traded goods are capital goods, a unit value index is constructed to proxy for the price of each of the above definitions of capital goods using United States import data and standard price index methodology. This index is then converted to Brazilian currency units using the real (R$)/US dollar ($) nominal exchange and multiplied by the Brazilian tariff rate to reflect the current domestic Brazilian price of these capital goods. We explored three types of indexes as well as two approaches to identify and delete outliers likely to be errors. The preferred index is a chain index with impact outliers removed at a 10 percent impact threshold. This section outlines in detail the construction of the capital price index, as well as other methodological issues considered.

## a. United States import data

Beginning in 1989, the Harmonized System (HS) of commodity classification has been used to classify imports and exports at a highly disaggregated level. The particular application of the HS system to United States imports is called the Harmonized Tariff Schedule (HTS). The Center for International Data at UC Davis makes available United States import data for 1989-2006 according to the 10-digit HTS number sourced from the United States Census Bureau and compiled by Feenstra, Romalis and Schott (2002). The import data is distinguished by source country and includes both quantitative information about imports and descriptive information about each commodity (see Feenstra, Romalis and Schott (2002) for documentation of the data).[23,24] Quantity, units of quantity, and customs value of general imports were extracted at the 10-digit HTS number and aggregated across the United States' trading partners for each year. The customs value reflects the value of imports as appraised by

[23] The quantitative information available includes: quantity of general imports (the total physical arrivals of merchandise from foreign countries, whether such merchandise enters consumption channels immediately or is entered into bonded warehouses or foreign trade zones under customs custody) and imports for consumption (the total of merchandise that has physically cleared through Customs either entering consumption channels immediately or entering after withdrawal for consumption from bonded warehouses under customs custody or from Foreign Trade Zones); customs value of general imports and imports for consumption; dutiable value (the customs value of foreign merchandise imported into the United States which is subject to duty); calculated duties (the estimated duty collected); and import charges for imports for consumption (equal to freight plus insurance).

[24] Feenstra (1996) also provides quantity, units of quantity, and customs value data for general imports for 1972-1988, reported at the 5-digit SITC Rev. 3 numbers. It is not possible to disaggregate this earlier data to the 10-digit HTS number to be appended to the 1989-2006 dataset (since a 5-digit SITC Rev. 3 number can be mapped to multiple 10-digit HTS numbers). Although it is possible to link the 10-digit HTS numbers to the 5-digit SITC Rev. 3 numbers, it is not possible to aggregate the value and quantity data from the 10-digit HTS number to the 5-digit SITC Rev. 3 number, since 10-digit HTS numbers grouped within the same 5-digit SITC Rev. 3 number are measured in different units. A possible solution is to calculate indices at the 10-digit HTS number for the 1988-2006 data, and then construct an index aggregated at the 5-digit SITC Rev. 3 numbers to be comparable with the 1972-1988 dataset. However, three complications arise. First, it is not possible to aggregate customs value and quantity data across countries at the 5-digit SITC Rev. 3 number in the 1972-1988 dataset, since imports from different countries are reported in different units of measurement at the same 5-digit SITC Rev. 3 number. This can be overcome by constructing an elementary level unit value index for each commodity classification at the country level, and then aggregating across countries to be comparable with the 1989-2006 dataset. Second, because there is no overlapping year of data in either of the datasets, it is not possible to link these two indexes. A conservative solution would be to assume no price change between 1988 and 1989. Lastly, what has currently prevented this index from being constructed is lack of tariff data at the 5-digit SITC Rev. 3 level for Brazil from 1972-1988.

the United States customs service. This value is generally defined as the price actually paid or payable for merchandise when sold for exportation to the United States, excluding United States import duties, freight, insurance and other charges incurred in bringing the merchandise to the United States, and is the value on which duties are expressed.

Concordance tables were then used to map the three definitions of capital goods to the 10-digit HTS number. The United States Department of Commerce (1992) provides a concordance of the Bureau of Economics' 34-industry code to the 3-digit 1987 SIC numbers. Concordance tables from the United Nations Statistics Division's Classification Registry were then used to map the 3-digit 1987 SIC numbers as well as the United Nations' 3-digit BEC numbers to the 5-digit SITC Rev. 3 numbers. Feenstra, Romalis and Schott (2002) provide concordances between the 10-digit HTS numbers and the 5-digit SITC Rev. 3 numbers in their data files. Of the 24,947 10-digit HTS numbers provided in Feenstra, Romalis and Schott's (2002) 1989-2006 data files, 2,730 were classified as a capital good under the BEC definition, 4,460 under the SITC7 definition, and 5,864 under the BEA definition.

However, further data editing was performed on the 1989-2006 dataset. First, observations were dropped due to missing quantity data despite a positive customs value reported (a zero value for quantity indicates that the units could not be measured). Second, an HTS number was dropped from the sample if different units of quantity were provided for the same 10-digit HTS number. Although there were no 10-digit HTS numbers for which different units of quantity were reported within a year, there were situations in which the unit of quantity changed over the years. Third, a 10-digit HTS number was dropped if the unit of quantity was not provided in any year, despite having both quantity and value data (and unit of quantity for other years). This was necessary to ensure that all units of quantity were consistently being measured over time as well as within a year. Fourth, classifications were dropped if unit values were observed for three years or less. Fifth, a classification was dropped if it was identified as a type of capital good in one year and not the next (there were five instances where the concordance in Feenstra, Romalis and Schott's (2002) data files mapped the same 10-digit HTS number to different 5-digit SITC Rev. 3 numbers in different years, and not all of these SITC Rev. 3 numbers were classified as capital goods). And sixth, a classification was dropped if tariff data was missing for any particular year. As a result of the data editing, 2,117 10-digit HTS numbers are capital goods under the BEC definition, 3,020 under the SITC7 definition, and 3,843 under the BEA definition over the 18 years.

It is also important to note that the HTS code reported in Feenstra, Romalis and Schott's (2002) data files is the HS code in use during the year in which the import data corresponds. The HS system has gone through multiple revisions since 1988, including 1996, 2002 and 2007. Because concordances do not exist at the 10-digit HTS number, the same commodity identified under different 10-digit HTS numbers in the different revisions cannot be identified.

## b. Unit value index

For each individual commodity class, unit values in any period are measured as the total value of shipments divided by the corresponding total quantity. Unit value indexes are the ratio of the unit value in the current period to the unit value in the reference period.[25] Elementary level unit value ratios, also referred to as elementary level unit value indexes, for each commodity class $i$, $I_i$, is a price comparison between the current period $t$ and a reference period 0 over $m = 1, \ldots, M$ items in period $t$ and over $n = 1, \ldots, N$ items in period 0. That is,

$$I_i^{0:t} = \left(\frac{\sum_{m=1}^{M} p_m^t q_m^t}{\sum_{m=1}^{M} q_m^t}\right) / \left(\frac{\sum_{n=1}^{N} p_n^0 q_n^0}{\sum_{n=1}^{N} q_n^0}\right)$$

where the prices and quantities are given, respectively, by $p_m^t$ and $q_m^t$ for period $t$, and $p_m^0$ and $q_m^0$ for the reference period 0. In practice, total customs value and total quantity of each commodity class $i$ at period $t$ and the reference period 0 is observed. Because unit value

[25] Note that three types of reference periods can be distinguished: weight reference period (the period covered by the expenditure statistics used to calculate the weights); price reference period (the period whose prices are used as denominators in the index calculation); and index reference periods (the period in which the index is set to 1). In the methodology of this study, the weight reference period, price reference period, and index reference period are the same.

indexes are used to represent price changes, the items within the commodity classes for which transactions are aggregated must be homogenous items.[26] Therefore, the commodity class is defined at the 10-digit HTS number, which is the most disaggregated customs value and quantity data.

Once the elementary level unit value indexes are constructed, they are subsequently aggregated across commodity classes using standard weighted index number formulas.[27] Different methods can be used to aggregate across commodity classes to construct the aggregated price index. We explore multiple methods, including a base index, a spliced index, and a chain index. Although the capital price deflator has been calculated using each of the above types of indexes for comparison, the chain index methodology is the preferred methodology adopted in this exercise.

### i. Base index

A standard base index constructs the elementary level unit value indexes by relating the unit value of commodity class $i$ in period $t$ to the unit value in the base period 0, then aggregates across weighted commodity classes,

$$I^{0:t} = \sum_i w_i^0 I_i^{0:t}.$$

In this exercise, 1989 is treated as the reference period such that $I^{1989} = 1$.

The major drawback of a base index is that each commodity class's unit value must be observed in the base period (or in each period if the weights are kept constant). Only about 40 percent of the 10-digit HTS numbers considered capital goods under any of the definitions had customs value and quantity data observed for all years 1989-2006 (43 percent for capital goods under the BEC definition, 40 percent under the SITC7 definition, and 42 percent under the BEA definition). A base index was calculated using only these commodity classes, but

[26] Otherwise bias will exist due to a compositional shift in the basket of heterogeneous transacted items.

[27] Export and import price indices have price changes of well defined representative items derived from establishment surveys at the elementary level. Import unit value indices differ from price indices because of their source data.

two alternative methods were considered that allow for inclusion of the other commodity classes with missing unit value observations.

### ii. Spliced index

A spliced index allows for the inclusion of a new commodity class into a base index. Because the unit value of this new commodity class was not observed in the base period, it is necessary to estimate its unit value in the base period using price ratios calculated for the items that remain, a subset of these items, or some other indicator. In practice, the base unit value for this commodity class becomes the unit value in the first year the new commodity class is included into the index, with its elementary level unit value index equaling 1 in that year. Its elementary level unit value index is then tracked according to that year, but is then subsequently deflated to the base year using the aggregated base index discussed above. For a commodity class that disappeared during the sample period, its weight is simply treated as 0.

### iii. Chain index

Another way to include new commodity classes into an index is to construct a chain index. A chain index is not only preferred when many new commodity classes are entering and disappearing, but a chain index also appropriately incorporates the introduction of new weights every year. In fact, an index constructed using fixed weights will be biased upwards due to a negative correlation between prices and quantities.

A chain index is constructed as multiple base indexes linked together. To link each of the base indexes, an overlapping period is needed in which the index is calculated using both the old and the new weights. Each of the base indexes is then multiplied to form a chain index with period 0 as the reference period. That is,

$$
\begin{aligned}
I^{0:t} &= \sum_i w_i^0 I_i^{0:1} \sum_i w_i^1 I_i^{1:2} \dots \sum_i w_i^{t-1} I_i^{t-1:t} \\
&= I^{0:1} I^{1:2} \dots I^{t-1:t}
\end{aligned}
$$

where $i$ represents the commodity class.

Given the nature of the United States import data at the 10-digit HTS number, a chained index methodology is preferred over a base index or spliced index methodology for primarily two reasons. First, rather than calculating a price index for a 'fixed basket of goods', which is what a base index is intended to capture, the exercise is to calculate a price index for 'capital goods'. Since new types of capital goods are continually being imported for the first time as well as exiting the market, new commodity classes are appearing and disappearing throughout the time frame. A chained index is better suited to allow for new commodity classes to be included in the index. This is further necessitated by the fact that the Feenstra, Romalis and Schott (2002) import data only provides the current HS code. So although the same product may be imported throughout the time frame, its 10-digit HTS number may change and is thus identified as a new commodity class. Without a concordance table at the 10-digit HTS level, it is impossible to identify and adjust for these changes. The second reason a chain index is preferred is because new weights are being introduced every period. When new weights are introduced, a new index should be calculated using the new weights, with an overlapping period in the old index using the old weights. The two indexes are then multiplied to form a chain index.

## c. Weights

The weights applied to the unit value changes should represent each elementary level unit value index's importance to the overall index. Thus an appropriate weighting scheme, $w$, is the share of the commodity class $i$'s customs value, $CV_i$, to the total value of imports for each year $t$,

$$w_i^t = \frac{CV_i^t}{\sum_i CV_i^t}.$$

This weighting scheme is consistent with price index theory, is also utilized by the Bureau of Labor Statistic's Import and Export Price Index, and is the methodology outlined in the

Export and Import Price Index Manual: Theory and Practice (International Monetary Fund 2009).

## d. Example of base, spliced, and chain unit value indexes

Table 2 provides an example of each of the different unit value index methodologies outlined above.

**Table 2: Example of Base, Spliced, and Chain Unit Value Indexes**

| **Commodity Class $i$** | $UV_i^0$ | $w_i^0$ | $UV_i^1$ | $w_i^1$ | $UV_i^2$ | $w_i^2$ | $UV_i^3$ | $w_i^3$ |
|---|---|---|---|---|---|---|---|---|
| A | 2 | 0.4 | 2.5 | 0.3 | 2.3 | 0.35 | 2.6 | 0.4 |
| B | 5 | 0.6 | 4.9 | 0.5 | 5.1 | 0.55 | 5.5 | 0.45 |
| C | | | 30 | 0.2 | 33 | 0.1 | 35 | 0.15 |
| **Commodity Class $i$** | $I_i^0$ | $w_i^0$ | $I_i^1$ | $w_i^1$ | $I_i^2$ | $w_i^2$ | $I_i^3$ | $w_i^3$ |
| **Base Index** | | | | | | | | |
| A | 1 | 0.4 | 2.5/2=1.25 | n.a. | 2.3/2=1.15 | n.a. | 2.6/2=1.3 | n.a. |
| B | 1 | 0.6 | 4.9/5=0.98 | n.a. | 5.1/5=1.02 | n.a. | 5.5/5=1.1 | n.a. |
| | **$I^0$=1** | | **$I^{0:1}$=0.4*1.25+0.6*0.98 =1.088** | | **$I^{0:2}$=0.4*1.15+0.6*1.02 =1.072** | | **$I^{0:3}$=0.4*1.3+0.6*1.1 =1.18** | |
| **Spliced Index** | | | | | | | | |
| A | 1 | 0.4 | 2.5/2=1.25 | n.a. | 2.3/2=1.15 | n.a. | 2.6/2=1.3 | n.a. |
| B | 1 | 0.6 | 4.9/5=0.98 | n.a. | 5.1/5=1.02 | n.a. | 5.5/5=1.1 | n.a. |
| Base Index | $I^0$=1 | | $I^{0:1}$ =1.088 | | $I^{0:2}$ =1.072 | | $I^{0:3}$ =1.18 | |
| C | | | 1 | 0.8 | 33/30=1.1 | n.a. | 35/30=1.1667 | n.a. |
| Spliced C | | | | 0.2 | =1.1*1.088 =1.1968 | n.a. | =1.1667*1.088 =1.2694 | n.a. |
| | **$I^0$=1** | | **$I^{0:1}$=1.088** | | **$I^{0:2}$=0.8*1.072+0.2*1.1968 =1.0970** | | **$I^{0:3}$=0.8*1.18+0.2*1.2694 =1.1979** | |
| **Chain Index** | | | | | | | | |
| A | 1 | 0.4 | 2.5/2=1.25 | 0.3 | 2.3/2=1.15 | 0.35 | 2.6/2.3=1.13 | n.a. |
| B | 1 | 0.6 | 4.9/5=0.98 | 0.5 | 5.1/5=1.02 | 0.55 | 5.5/5.1=1.08 | n.a. |
| C | | | 1 | 0.2 | 33/30=1.1 | 0.1 | 35/33=1.17 | n.a. |
| Base Index | $I^0$=1 | | $I^{0:1}$=0.4*1.25+0.6*0.98 =1.1088 | | $I^{1:2}$=0.3*1.15+0.5*1.02+0.2*1.1 =1.075 | | $I^{2:3}$=0.35*1.13+0.55*1.08+0.1*1.17 =1.1065 | |
| | **$I^0$=1** | | **$I^{0:1}$=1*1.088 =1.088** | | **$I^{0:2}$=1.088*1.075 =1.1696** | | **$I^{0:3}$=1.1696*1.1065 =1.2942** | |

Notes: Table 2 provides an example of each of the different unit value index methodologies outlined above, including a base index, a spliced index, and a chain index. The first three rows give, for each capital good (commodity class) A, B and C, their respective unit value and weight in four time periods.

## e. Outliers

Data editing is necessary to detect outliers and correct for possible errors. Outliers are identified as observations that fall outside some pre-specified acceptance interval and are judged to be unrealistic. Ideally, one could detect outliers and then verify and correct any errors, such as recording or coding mistakes. However, given the nature of the dataset, outliers are treated as errors and deleted. Two different methods were used to identify and delete outliers, adopted from the Import and Export Price Index Manual: Theory and Practice (International Monetary Fund 2009).

First, identification of unusual unit value changes was conducted through statistical checking of input data. Statistical checking of input data compares, for some time period, each unit value's change to the entire distribution of unit value changes. Kernel density plots for each index show that the distribution of unit value changes over all years for each definition of capital is skewed significantly left (see appendix b). As such, the 10-digits HTS numbers causing the unit value changes at the top 2.5 percent of the distribution were dropped and the indexes were recalculated without these products. Although it is necessary to drop the commodity for all years for the base index (otherwise the sample would no longer be constant), it is not necessary to drop the commodity for all years for the splice and chain indexes, merely the year of the unusual unit value change. This is another reason a chain index is preferred

Second, identification of unusual unit value changes was conducted through output checking, also called checking by impact. This procedure is based on calculating the impact an individual unit value change has on the index to which it contributes, measured as the unit value's weight multiplied by the price relative and divided by the level of the index to which it contributes. That is, the impact on the overall index $I$ of the change of the unit value of commodity class $i$ from the base period 0 to period $t$ is $\frac{w_i^0(\frac{I_i^t}{I_i^0})}{I^0}$. A minimum threshold for this impact was set, such that any 10-digits HTS numbers that had a price change that caused an impact greater than this threshold was dropped and the indexes recalculated without these products. Table 3 reports the percentage of observations dropped from each type of index for the different definitions of capital under various minimum thresholds.

**Table 3: Percentage of Observations Deleted under Different Impact Thresholds**

| Impact Threshold | Base Index | Spliced Index | Chain Index |
|---|---|---|---|
| **BEC** | | | |
| 20% | 0.11% | 0.06% | 0.09% |
| 10% | 0.33% | 0.11% | 0.09% |
| 5% | 0.44% | 0.27% | 0.39% |
| 1% | 3.40% | 1.37% | 2.00% |
| **SITC7** | | | |
| 20% | 0.08% | 0.03% | 0.05% |
| 10% | 0.08% | 0.03% | 0.07% |
| 5% | 0.33% | 0.08% | 0.10% |
| 1% | 2.17% | 0.62% | 1.47% |
| **BEA** | | | |
| 20% | 0.06% | 0.03% | 0.04% |
| 10% | 0.06% | 0.03% | 0.05% |
| 5% | 0.18% | 0.09% | 0.10% |
| 1% | 1.53% | 0.80% | 0.94% |

Notes: Table 3 presents, for each definition of capital goods (BEC, SITC7, BEA), the percentage of observations deleted when using an impact threshold approach to identify outliers for different thresholds (1%, 5%, 10%, 20%). Each column represents a different unit value index methodology (base, spliced, chain).

Output checking is the adopted methodology to detect and delete outliers because it identifies the unit value changes likely to be errors that have significant impacts on the overall index without losing significant amounts of information. A minimum impact threshold of 10 percent was adopted. Significant differences exist between the indexes using a 20 percent threshold level and a 10 percent threshold level, but few differences are made adopting a more stringent threshold. Furthermore, outliers were detected before the indexes were converted to Brazilian prices. For the chain index, impact outliers were detected by their impact on each year's base index, that is, before each of the base indexes were multiplied to form the chain index. The calculation of each type of index using the different methods to detect outliers is presented below and appendix c presents the preferred index calculated using different minimum thresholds.

## f. Brazilian nominal exchange rate

Once the elementary level unit value indexes are constructed at the 10-digit HTS number, which captures the price change of capital goods imports in nominal US dollars, the price indexes can be converted into Brazilian currency units by multiplying by the real(R$)/US dollar($) nominal exchange rate, sourced as the Office Exchange Rate (LCU per US$, period average) from the World Bank's World Development Indicators, defined as: "Official exchange rate refers to the exchange rate determined by national authorities or to the rate determined in the legally sanctioned exchange market. It is calculated as an annual average based on monthly averages (local currency units relative to the U.S. dollar)."

Rebasing is then necessary such that the elementary level unit value indexes equal 1 in the reference period. To rebase a base or spliced index, it is necessary to divide the index by the exchange rate in the base year 0, since

$$\frac{\frac{I_i^t}{I_i^0}e^t}{\frac{I_i^0}{I_i^0}e^0} = \frac{I_i^t e^t}{I_i^0 e^0}$$

where $e$ is the real/US dollar nominal exchange rate. To rebase a chain index, since the base year is the previous year, it is necessary to divide the index by the exchange rate in the previous year, since

$$\frac{\frac{I_i^t}{I_i^{t-1}}e^t}{\frac{I_i^{t-1}}{I_i^{t-1}}e^{t-1}} = \frac{\frac{I_i^t}{I_i^{t-1}}e^{t-1}}{e^{t-1}} = \frac{I_i^t e^t}{I_i^{t-1} e^{t-1}}.$$

## g. Brazilian tariff data

To accurately reflect the current domestic Brazilian price of the capital goods imports, the elementary level unit value indexes are multiplied by the Brazilian tariff rate,

$$I_i^t = I_i^t(1+\tau)$$

where $\tau$ is the simple average ad valorem tariff rate that Brazil applies to the world for commodity class $i$. The index is then rebased appropriately as outlined above.

The entire tariff structure is available for bulk download at the 6-digit HTS number from the TRAINS database, managed by UNCTAD, available through the World Bank's World Integrated Trade Solutions (WITS) database for the time period considered. Because the elementary level unit value indexes are constructed at the 10-digit HTS number, it is necessary to assume that all tariff lines under the sub heading of the 6-digit HTS number are charged the same MFN applied rate. Furthermore, collecting tariff data from a Brazilian source at a level lower than the 6-digit HTS number is inaccurate, since the data is United States imports, and only the 6-digit HTS number is consistently used internationally.[28] However, only 0.02 percent of the observations at the 10-digit HTS number had no tariff data reported for the BEC definition of capital, 0.01 percent for the SITC7 definition, and 0.08 percent for the BEA definition. These observations were dropped from the sample.

TRAINS reports the applied MFN tariff rate, the effectively applied tariff rate, and the bound duty rate. The bound duty rate is inappropriate as many countries apply a tariff rate that is lower than their negotiated duty rates. The effectively applied rate is the lowest available tariff. It is the MFN applied rate unless another tariff is recorded for the considered product (mostly accounting for the fact that some countries are given preferential treatment, although a country may apply a tariff higher than the MFN rate to non-World Trade Organization members). The MFN applied rate is chosen over the effectively applied tariff rate, justified by the Brazilian trade reforms occurring during this time period, discussed below.

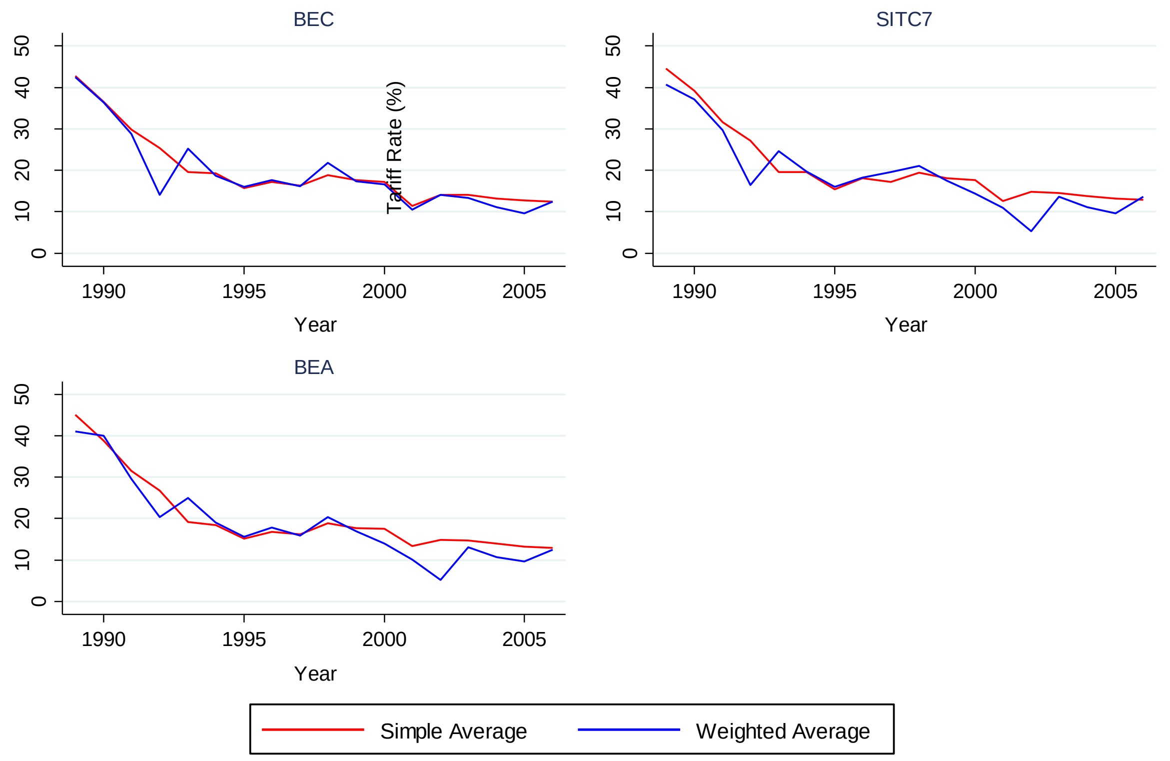
TRAINS also reports both a simple average tariff rate as well as a trade-weighted tariff rate for Brazilian imports from the world. The simple average tariff rate was used rather than a trade-weighted tariff rate as to not bias the capital price index. First, the trade-weighted tariff rate would include another weight in front of $\tau$. This is inappropriate since the index $I_i^t$ is subsequently multiplied by a more appropriate trade weight determined by world trade when aggregating to the overall index.[29] Second, the weight applied to the trade-weighted tariff is a distorted weight in itself as it endogenously underestimates the level of protection since imports are inversely related to the tariff rate. However, there are little differences in the averages of the MFN applied simple-average tariff rate and the MFN applied trade-weighted

[28] Using Brazilian reported applied tariff rates for 2011 from Receita Federal (2011), 92 percent of tariff lines have the same rate applied to all 8-digit HTS numbers within the 6-digit HTS number, and 96.3 percent have the same rate applied at all the 10-digit HTS numbers within the 8-digit HTS number.

[29] It is noted that the weights used to construct the index are the United States' import weights from the world, not Brazil's as is captured in the trade-weighted tariff.

tariff rate reported by TRAINS at the 6-digit HTS number for each of the definitions of capital goods, as Figure 1 illustrates.

**Figure 1: Simple versus Weighted Average Tariff**




Notes: Figure 1 plots Brazil's simple average (red line) and weighted average (blue line) tariff on imported capital goods from 1989-2006 for each definition of capital goods (BEC, SITC7, BEA). The simple average and weighted average tariffs were calculated using reported tariff rates of Brazil for the world at each 6-digit HTS number, then aggregated accordingly.

## h. Brazilian trade reforms

Beginning in the 1950s, Brazil implemented restrictive trade policies as a means of promoting industrialization development strategies, heightening by the mid 1960s. Between 1967 and 1973, during Brazil's period of significant economic growth, Brazil adopted a relatively more open trade policy in comparison to the previous decade. However, Brazil reversed this trend in 1973 in response to the steep rise in world oil prices by again restricting imports and increasing tariffs.

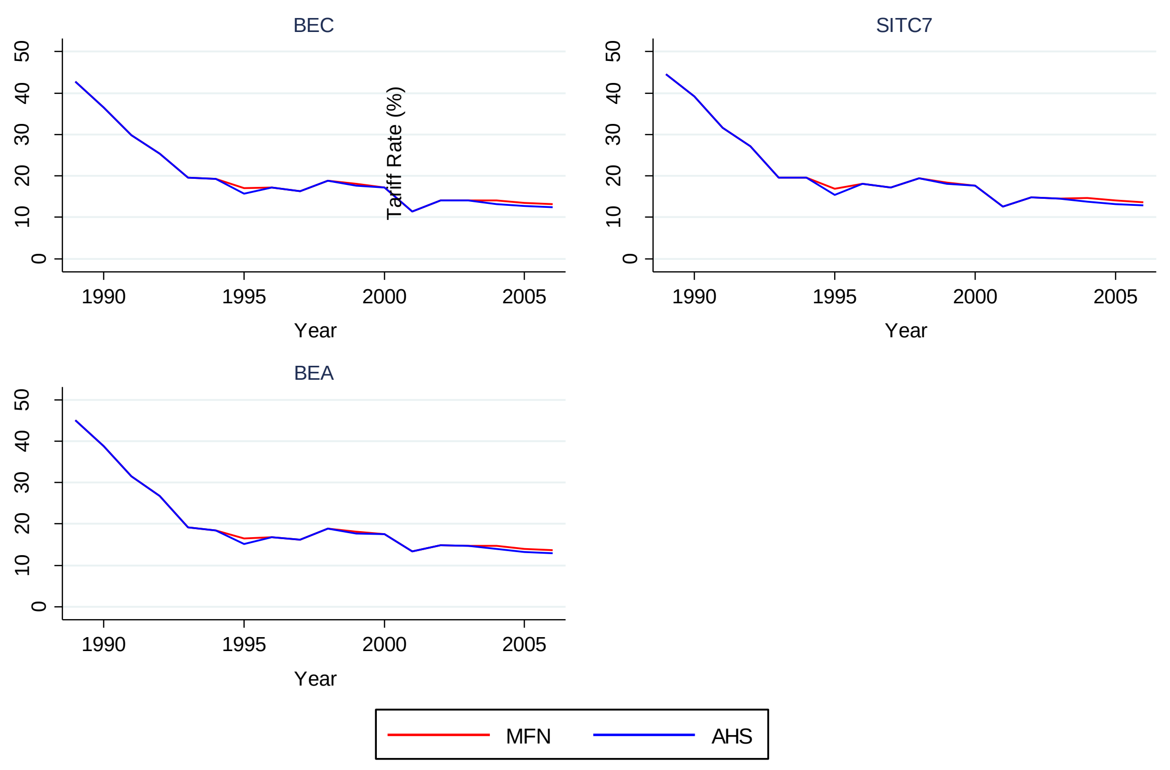
Only in the mid-1990s did Brazil begin to significantly liberalize its trade policy, phasing in a series of tariff reductions in the early part of the decade. MERCOSUR was founded in 1991 and updated in 1994, which allowed free trade for most goods between Brazil, Argentina, Paraguay and Uruguay. Associate member status was later extended to Chile, Ecuador, Bolivia, Colombia and Peru.

Brazil has signed other regional trade agreements: Protocol on Trade Negotiations (PTN) in 1973, Global System of Trade Preferences among Developing Countries (GSTP) in 1989, Latin American Integration Association (LAIA) in 1981, and MERCOSUR-India in 2009 (World Trade Organization 2011). These regional trade agreements are signed with other Latin American countries as well as other developing countries in Africa and Asia. Brazil does not extend preferential treatment to developed countries, including the US, Europe and Japan.

Thus, for the purposes of this study, the MFN applied tariff rate is preferred over the effectively applied tariff rate for two reasons. First, given the nature of Brazil's preferential arrangements during the period 1989-2006, it is unlikely that the MERCOSUR countries and the other countries to which Brazil extends preferential treatment are sources of capital goods imports. Most of the world's capital goods are provided by a small number of research and development intensive countries, all of which would face the applied MFN rate. And most countries, in particular developing countries, tend to import a large fraction of their capital goods. In fact, purchases from France, Germany, Italy, Japan, Sweden, the United Kingdom and the United States account for 70 percent of these foreign purchases (Eaton and Kortum 2001; Alfaro and Ahmed 2007). Second, when considering the tariff rate that Brazil extends to the world, the effectively applied rate must have an additional weighting system where the reported tariff is equal to the weighted average of the different tariffs charged to different countries.

Thus, when comparing the effectively applied rate to the applied MFN rate for capital goods, it is expected that these two series should be very close. This can be seen in Figure 2, which plots the average applied MFN rate against the average effectively applied rate for each year. The two series are identical for most years, however in the years that they do diverge, it is clear that the average effectively applied rate is lower than the average applied MFN rate. Furthermore, Table 4 reports the correlation coefficient between the MFN applied and effectively applied tariff rates. As expected, strong correlations exist between the two series for each of the definitions of capital goods, both for all years and within each year.

**Figure 2: Simple Average MFN versus AHS Tariff**




Notes: Figure 2 plots Brazil's simple average most-favored nation (red line) and simple average effectively applied (blue line) tariff on imported capital goods from 1989-2006 for each definition of capital goods (BEC, SITC7, BEA). The MFN is the tariff rate that Brazil applies to WTO members. The effectively applied is the tariff rate that Brazil actually applies, considering Brazil grants preferential access to some trading partners. The simple average tariffs were calculated using reported tariff rates of Brazil for the world at each 6-digit HTS number, then aggregated accordingly.

**Table 4: Correlation Coefficient between MFN and effectively applied tariff rates**

| Year | BEC | SITC7 | BEA |
|---|---|---|---|
| All years | 0.9869 | 0.9896 | 0.9924 |
| 1989 | 1 | 1 | 1 |
| 1990 | 1 | 1 | 1 |
| 1991 | 1 | 1 | 1 |
| 1992 | 1 | 1 | 1 |
| 1993 | 1 | 1 | 1 |
| 1994 | 1 | 1 | 1 |
| 1995 | 0.9855 | 0.9850 | 0.9877 |
| 1996 | 1 | 1 | 1 |
| 1997 | 1 | 1 | 1 |
| 1998 | 1 | 1 | 1 |
| 1999 | 0.9770 | 0.9859 | 0.9911 |
| 2000 | 1 | 1 | 1 |
| 2001 | 1 | 1 | 1 |
| 2002 | 1 | 1 | 1 |
| 2003 | 1 | 1 | 1 |
| 2004 | 0.8569 | 0.9158 | 0.9592 |
| 2005 | 0.9109 | 0.9445 | 0.9754 |
| 2006 | 0.9701 | 0.9822 | 0.9819 |

Notes: Table 4 presents the correlation coefficient between the most-favored nation and effectively applied tariff on imported capital goods from 1989-2006 for each definition of capital goods (BEC, SITC7, BEA). The MFN is the tariff rate that Brazil applies to WTO members. The effectively applied is the tariff rate that Brazil actually applies, considering Brazil grants preferential access to some trading partners.

### i. Adjusting for quality change

The measurement of price changes is also complicated by changes in the quality of existing goods. Over time, the quality of what is produced changes, such that observed changes in prices may arise partly from quality changes. Ideally, adjustments would to be made such that the index is not capturing price changes due to quality changes. Then the capital stock series could be constructed after deflating nominal investment flows by a quality-adjusted price index. Otherwise, the growth of the capital stock and its estimated contribution to output growth may be biased downward.

This analysis ignores quality differences over time, and the unit value of commodity class $i$ in period $t$ is compared to that of the reference period assuming no quality changes (or put another way, that any change in quality has no affect on price). Although different methods do exist to adjust prices for quality differences, these methods are not implementable due to the nature of the import data.[30]

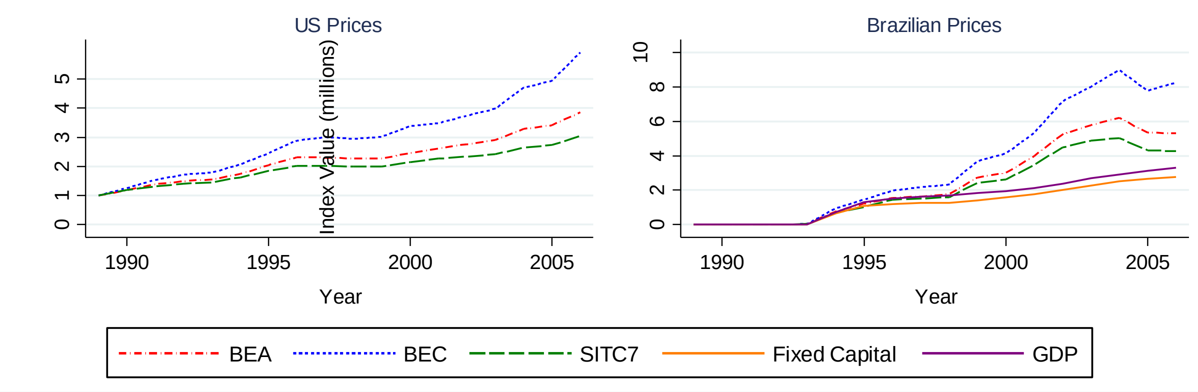
## 5. Capital Price Index Results

Here we present the preferred capital price index for each of the definitions of capital goods in nominal United States prices as well as nominal Brazilian prices. The indexes in Brazilian prices are also plotted against Brazil's fixed capital formation price deflator and Brazil's GDP

[30] For example, overlap pricing. If a comparable replacement can be made of an item with a quality upgrade, and if observed prices overlap for the same period of the old version and the new version, a quality adjustment can be made such that the price difference between the old item and its replacement reflects the effect of the quality difference on price. It is also possible to estimate the effect of the quality change on prices for non-comparable replacements, for example, using hedonic regressions. All price-determining characteristics are recorded for commodities and statistical techniques are used to estimate the implicit prices of product characteristics. These implicit prices help disaggregate the observed price difference between two products into quality change and pure price change. See Chapter 7 of Consumer Price Index Manual: Theory and Practice (International Monetary Fund 2004) for a more detailed discussion of quality adjustment techniques. Also see Sakellaris and Vijselaar (2004) who derive quality-adjusted price indices using United States data to construct an appropriately deflated capital series to re-calculate TFP in the euro area.

deflator downloaded from the Brazilian Institute of Geography and Statistics as to highlight the significant divergence in price deflators.

**Figure 3: Capital Price Index, 1989-2006**



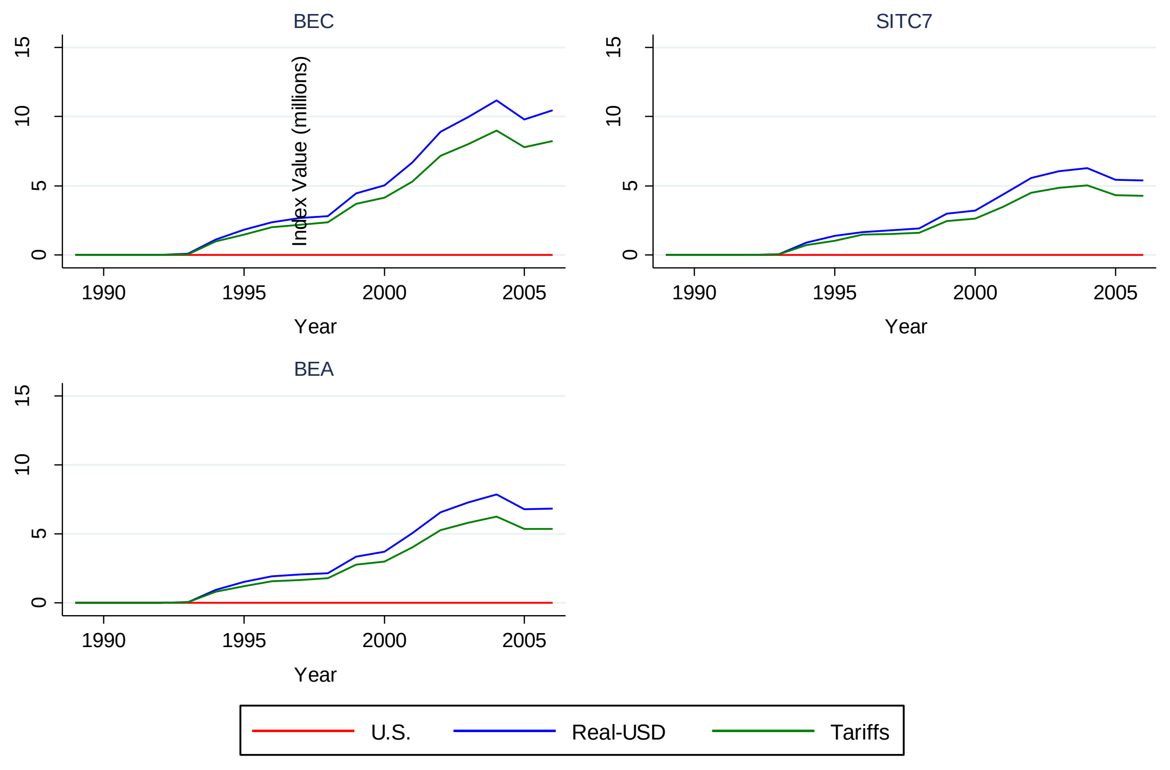

Notes: Figure 3 plots the capital price index in US prices (left panel) and Brazilian prices (right panel) for the three definitions of capital goods (BEA in red, BEC in blue, SITC7 in green) and Brazil's published price deflators (fixed capital in orange, GDP in purple) from 1989-2006. The index is based to equal 1 in the base year (1989).

Figure 3 shows increases in each of the capital price series measured in United States prices between 1989 and 2006, with cumulative increases of 440 percent for capital according to the BEC definition, 190 percent for the SITC7 definition, and 263 percent for the BEA definition. Once converted to Brazilian prices, the inflation of the price of capita is huge, with cumulative increases of 652 million for the BEC definition, 308 million percent for the SITC7 definition, and 402 million percent for the BEA definition. This massive increase is driven by the nominal exchange rate due to depreciation of the real against the US dollar during a period of hyperinflation in Brazil. More importantly is the significant divergence between each of the capital price indexes and Brazil's fixed capital formation price deflator and GDP deflator. Each of the constructed capital price series lies above Brazil's fixed capital formation price deflator and GDP deflator. In 2006, the percentage difference between the GDP deflator and the capital price index according to the BEC definition is 128 percent, the SITC7 definition is 23 percent, and the BEA definition is 52 percent. For the fixed capital formation price deflator, these numbers are 173 percent, 48 percent, and 82 percent, respectively.

These results substantiate the argument of this study: Brazil's official price deflators are underestimated, and this divergence between the price of capital and the published Brazilian price deflators causes mis-measurement of Brazil's TFP estimates. Furthermore, it is not simply an inflation story. Inflation would affect both the GDP deflator and the price of capital and thus the relative prices would not change. What matters is that the price of capital rose more than the price of GDP.

It is possible to decompose how much of the growth in the price of capital is due to the evolution of the global price of capital goods, the nominal exchange rate, and the changing tariff structures. These results are presented in Figure 4 for each of the definitions of capital goods. As can be seen, the inflation of the price of capital is driven by the nominal exchange rate due to depreciation of the real against the US dollar. However, the capital price index falls after taking into account Brazil's declining tariffs on capital goods during the trade reforms of this period. (See sections 4.7 and 4.8 for a discussion of Brazilian tariff rates and Brazil's trade reforms.) This "trade policy effect", however, is not enough to outweigh the "capital stock inflation effect".

**Figure 4: Decomposing the Growth of the Capital Price Index**

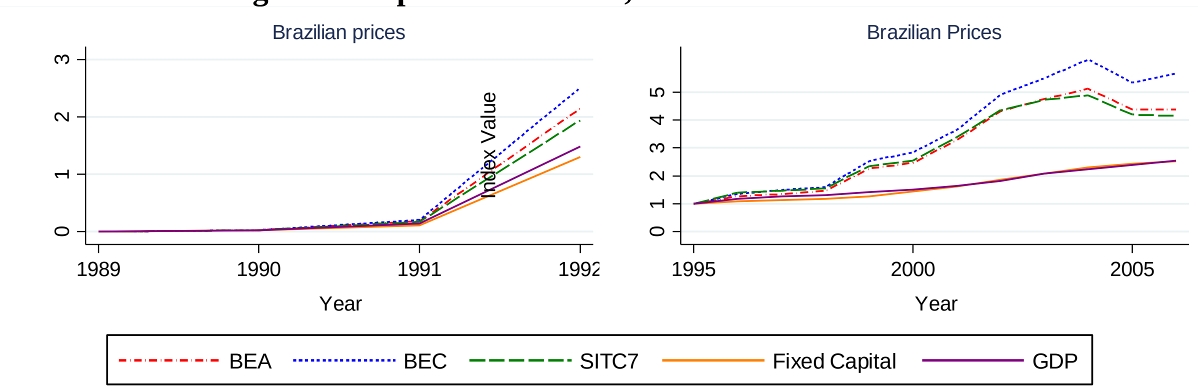



Notes: Figure 4 plots the capital price index for the three definitions of capital goods (BEA, BEC, SITC7) first in US prices (red), second in Brazilian prices excluding tariffs (blue), and third in Brazilian prices including tariffs (green) from 1989-2006. The index is based to equal 1 in the base year (1989).

The results illustrate a clear divergence between the capital price indexes and Brazil's fixed capital formation price deflator and GDP deflator in the mid-1990s. Therefore we expect a stronger recovery in TFP after 1992 than has previously been observed. However, whether divergence is occurring in the lower part of the series is also important. The TFP estimates for Brazil documented by Gomes, Pessôa and Veloso (2003) show declining TFP levels beginning in 1982, continuing until 1992, and then stabilizing after 1992. Zooming in on the lower part of the capital price index shows that only in 1991 do the constructed capital price indexes begin to diverge from Brazil's fixed capital formation price deflator and GDP deflator (Figure 5). Therefore, it is expected that the new TFP series will still decline prior to 1992.

Furthermore, the results also illustrate that the inflation of the price of capital is driven by the nominal exchange rate, which is due to the hyperinflation occurring in Brazil during the beginning of this time period. However, the real stabilization plan was implemented in 1994 and by 1995 inflation had dropped to double digits. Therefore, we explore whether the results still hold starting in 1995 after stabilization of the real (Figure 5). Again, each of the constructed capital price series lies above Brazil's fixed capital formation price deflator and GDP deflator, again supporting the argument that it is the divergence in the relative prices that drive the results, not simply hyperinflation.

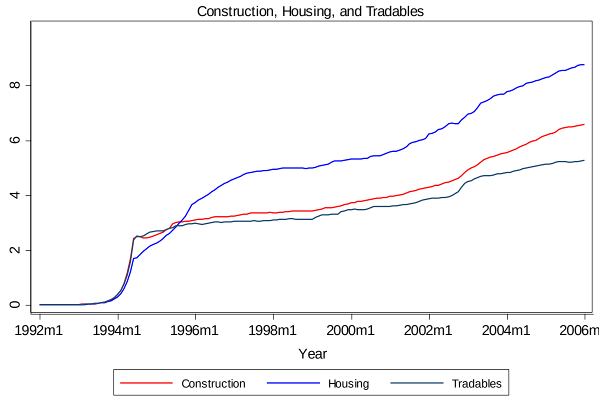
**Figure 5: Capital Price Index, 1989-1992 and 1995-2006**




Notes: Figure 5 plots the capital price index in Brazilian for the three definitions of capital goods (BEA in red, BEC in blue, SITC7 in green) and Brazil's published price deflators (fixed capital in orange, GDP in purple) for two different time periods. In the left panel, the index is based to equal 1 in the base year 1989. In the right panel, the index is based to equal 1 in the base year 1995.

The standard practice for measuring capital accumulation is to take into account expenditures in structures. One drawback to the constructed capital price index is that we fail to capture the price of non-tradable capital goods, and therefore we lack a proxy for the price of capital that takes structures into account. Figure 6 plots Brazilian price indexes for construction, housing, and tradables in Brazil in 1992 prices. Tradables experienced a large price increase between 1994 and 1996 but construction and housing experienced larger price increases. This graph suggest that we are undervaluing the capital price index by excluding non-tradable capital investments such as structures, and thus any TFP estimates should be considered lower bounds.

**Figure 6: Brazilian Price Indexes (1992=1): Construction, Housing, and Tradables**




Source: Banco Central do Brasil (2012).
Notes: Figure 6 plots the price index of non-tradable goods (construction in red and housing in blue) and tradable goods (in green) in Brazilian prices from 1992-2006. The index values were calculated as a chain index using published month-on-month inflation series. The index is based to equal 1 in the base month (January 1992).

# 6. Brazilian Total Factor Productivity: 1989-2006

## a. Methodology

This section outlines the methodology used to calculate Brazil's TFP. First it is necessary to calculate a series of capital stocks from an investment series deflated by the capital price index. The PIM is the standard approach used in the literature. However, this is complicated by an appropriate assumption for the initial capital stock. Once the capital stock series is in hand, a series of TFP is calculated for Brazil using approaches similar to both Bugarin et al. (2002) and Gomes, Pessôa and Veloso (2003), as discussed above.

### i. Calculating the capital stock

Different methods exist to calculate a series of capital stocks. King and Levine (1994) assume that a country is continually at a steady state with a constant capital-output ratio, which implies that the rate of growth of the capital stock is equal to the rate of growth of output. Then the steady-state capital-output ratio is

$$\kappa = \frac{i}{\delta + g}$$

where $i$ is the steady state investment rate, $\delta$ is the depreciation rate of the capital stock, and $g$ is the growth rate of output.[31] Then the capital stock is calculated by observing a series of GDP values. However, this approach is only valid if the GDP deflator is equal to the price of capital, and is not an appropriate methodology to use for this study.

A second and more common approach is the PIM to calculate a series of capital stocks. The PIM assumes that the stock of capital is the accumulation of the stream of past investments,

$$K_{t+1} = (1 - \delta)K_t + I_t$$

where $I_t$ is the investment in period $t$, $K_t$ is the current period's capital stock, $K_{t+1}$ is the next period's capital stock, and $\delta$ is the depreciation rate of the capital stock. Once an investment series has been collected, the capital stock can then be calculated.

One challenge with the PIM is the need to assume a value of the initial capital stock. Nehru and Dhareshwar (1993) offer an overview of various ways to calculate the initial capital stock when using the PIM. (Appendix d shows how the results change when experimenting with different values of the initial capital stock.)

[31] Growth rates are computed as geometric averages. For example, the geometric average growth rate of output, Y, between date $t$ and $t + T$ is $g_{0,T} = \left(\frac{Y_T}{Y_0}\right)^{1/T} - 1$.

One possibility is to assume that the initial capital stock is zero. However, the capital series is sensitive to the initial capital stock value, although this sensitivity diminishes as the series progresses. Because the investment series used to construct the capital stock should be in real values, in practice, the first year that the investment series can be used is in 1989 since that is the base year of the capital goods price deflator. Thus obtaining a better estimate of the initial capital stock is important for this exercise.[32]

Harberger (1978) suggests a steady-state estimate of the initial capital stock, and is the most common approach adopted in the literature. To convert a flow variable (investment) to a stock variable (capital) in the steady state, it is necessary to divide the flow variable by the steady state growth in the flow variable – proxied by the average growth rate of the deflated investment series – and the depreciation rate (see, for example, Hall and Jones (1999)). However, after deflating the investment series with the capital price index, the average growth rate of the investment series is negative, both for the sample period 1989-2006 as well as the initial sample period, 1989-1993, which resulted in a negative capital stock. Nehru and Dhareshwar (1993) also ran into this problem, and explored an alternative way of obtaining the value of investment in the first period, which was to estimate a linear regression of the log of investment against time and calculate the fitted value of the initial investment level by this equation. This fitted investment level in the first period is then used to calculate the initial capital stock using the equation above. This could be done as an extension.

Rather, by assuming that the capital-output ratio is constant as in the steady state, the rate of growth of capital and output are theoretically equal during that period. Therefore the initial capital stock is calculated as

$$K_1 = \frac{I_1}{g + \delta}$$

[32] Both Bugarin et al. (2002) and Gomes, Pessôa and Veloso (2003) use the capital series from the Institute for Applied Economic Research (IPEA) for Brazil, and thus do not estimate an initial value of the capital stock. However, Gomes, Pessôa and Veloso (2003) estimate the initial capital stock for other countries in their study using the growth rate of the population and technological progress, as discussed in appendix d.

where $g$ is the growth rate of output and $I_1$ is the first year of the investment series.[33,34]

### ii. Calculating total factor productivity

To identify the contribution of productivity to the performance of the Brazilian economy, this study employs a growth accounting exercise to measure the Solow residual or TFP. As productivity cannot be measured directly, this approach measures TFP as a residual that accounts for output not caused by traditionally measured inputs of labor and capital. If inputs are accounted for, then TFP can be taken as a measure of an economy's long-term technological change or dynamism. Of course, the evolution of TFP is dependent on the specification of the production function that is chosen. As such, we use various specifications of the production functions to calculate TFP, but all have the properties of a neoclassical production function (homogeneity of first degree, positive marginal productivity, and decreasing returns to inputs).

First, TFP is calculated using a Hicks-neutral Cobb-Douglas production function of the form

$$y_t = A_t k_t^{\alpha}$$

where $y_t$ is output per worker at time $t$, $k_t$ is capital per worker, $\alpha$ is capital's factor share, and $A_t$ is TFP. This specification is the most basic approach that follows the methodology of Bugarin et al. (2002) as applied to Brazil. Thus, TFP can be calculated as

$$A_t = \frac{y_t}{k_t^{\alpha}}.$$

[33] As a benchmark, the average growth rate of output was calculated for the sample period 1989-2006. The growth rate was also calculated for the initial sample period, 1989-1993. However, during the sample time period, significant capital deepening is observed, resulting in an increasing capital-output ratio (albeit, likely because the capital stock is mis-measured). As such, the growth rate of output was also calculated for the period 1967-1976 which represents a time period with a constant capital output ratio found by Gomes, Pessôa and Veloso (2003) and may better approximate a steady state estimate of the growth rate of output. For robustness, we present the results using this approach in appendix d.

[34] Harberger (1978) suggests using an average growth rate of output as well as the corresponding average investment level, rather than the investment level in the initial year of the investment series, due to short-term variations in output and investment. Then, the calculated capital stock is centered in the middle of the period, and the recursion formula for capital accumulation would have to be applied in reverse to arrive at the initial capital stock. This was not done in this study.

Alternatively, country-specific or detrended TFP can be calculated after accounting for TFP at the technological frontier. Under this approach, aggregate output per worker can be represented by the following function

$$y_t = A_t k_t^{\alpha} (H_t \lambda_t)^{1-\alpha}$$

where $y_t$ is output per worker at time $t$, $A_t$ is country-specific or detrended productivity PTFD, $k_t$ is capital per worker, $\alpha$ is capital's factor share, $H_t$ is human capital (education) per worker, and $\lambda_t = (1+\gamma)^t$ represents productivity at the technological frontier in year $t$ where $\gamma$ is the rate of technological progress. Thus TFP is divided into two parts: PTFD given by $A_t$, which is specific to Brazil, and the contribution of the evolution of the technological frontier to productivity, ${\lambda_t}^{1-\alpha}$, which is common to all economies. This approach follows Gomes, Pessôa, and Veloso (2003) as applied to Brazil. The authors assume that education affects labor productivity according to the mincerian approach as incorporated into the literature of economic growth,

$$H_t = e^{\phi(h_{it})}$$

where $h_{it}$ denotes the average years of schooling of the labor force. Based on a decreasing relationship between average schooling and the rate of return to schooling, as observed in a cross-section of countries in various stage of development, the function $\phi(\cdot)$ is concave. In particular, the authors adopted the functional form

$$\phi(h_t) = \frac{\theta}{1-\psi} h_t^{1-\psi}.$$

where $\theta > 0$ and $0 < \psi < 1$. Therefore, PTFD for each year, given by $A_t$, is calculated as

$$A_t = \frac{y_t}{k_t^{\alpha}(H_t \lambda_t)^{1-\alpha}}.$$

In addition, we also calculate TFP using a constant elasticity of substitution (CES) production function, which relaxes the assumption of unitary elasticity of substitution between capital and labor and rather assumes a constant elasticity of substitution between capital and labor.[35] In other words, the production technology exhibits a constant percent change in factor proportions due to a percentage change in the marginal rate of technical substitution. We assume the production function takes the form

$$y_t = A_t[a(k_t)^{\frac{\sigma-1}{\sigma}} + (1-a)]^{\frac{\sigma}{\sigma-1}}$$

where $y_t$ is output per worker at time t, $k_t$ is capital per worker, $\alpha$ is capital's factor share, $A_t$ is TFP, $\sigma$ is the elasticity of substitution between capital and labor, and $0 < a < 1$ is the share parameter. This approach follows Eller-Jr. (2012) as applied to Brazil. Therefore, TFP can be calculated as

$$A_t = \frac{y_t}{[a(k_t)^{\frac{\sigma-1}{\sigma}}+(1-a)]^{\frac{\sigma}{\sigma-1}}}.$$

## b. Data

Brazil's labor force data was downloaded from the World Bank's World Development Indicators: Labor force (total), defined as "Total labor force comprises people ages 15 and older who meet the International Labour Organization definition of the economically active population: all people who supply labor for the production of goods and services during a specified period. It includes both the employed and the unemployed. While national practices vary in the treatment of such groups as the armed forces and seasonal or part-time workers, in general the labor force includes the armed forces, the unemployed, and first-time job-seekers, but excludes homemakers and other unpaid caregivers and workers in the informal sector."[36]

Brazil's nominal GDP series was downloaded from the World Bank's World Development Indicators: GDP (current LCU), defined as "GDP at purchaser's prices is the sum of gross value added by all resident producers in the economy plus any product taxes and minus any

[35] In addition, it is possible to impose factor-augmenting technological progress where the technology is not neutral between factors. First order conditions for profit maximization under perfect competition can be used to solve for labor versus capital productivity gains.
[36] Gomes, Pessôa and Veloso (2003) estimated the number of workers in Brazil using the active work force from Brazilian census data available every 10 years. The number was interpolated based on the growth rate between the 10-year intervals.

subsidies not included in the value of the products. It is calculated without making deductions for depreciation of fabricated assets or for depletion and degradation of natural resources. Data are in current local currency." The nominal GDP series was deflated using the GDP deflator downloaded from the World Bank's World Development Indicators: defined as "The GDP implicit deflator is the ratio of GDP in current local currency to GDP in constant local currency. The base year varies by country."

Brazil's average years of schooling was downloaded from Barro and Lee (2010): Average years of schooling for total population age group 15+. Data is available every five years from 1960-2010. For years not available, the number was interpolated based on the growth rate between the five-year intervals.

Brazil's investment series was proxied with Brazil's nominal gross fixed capital formation series, downloaded from the World Bank's World Development Indicators: Gross fixed capital formation (current LCU), defined as "Gross fixed capital formation (formerly gross domestic fixed investment) includes land improvements (fences, ditches, drains, and so on); plant, machinery, and equipment purchases; and the construction of roads, railways, and the like, including schools, offices, hospitals, private residential dwellings, and commercial and industrial buildings. According to the 1993 SNA, net acquisitions of valuables are also considered capital formation. Data are in current local currency." We acknowledge that this definition includes residential investments, which are subsequently included in the TFP estimates. While ideally we would remove these investments since these are not productive, their inclusion places an upward bias on the accumulation of capital and a downward bias on the TFP estimates.

Other parameter values were taken from Gomes, Pessôa, and Veloso (2003), including: $\alpha = 0.4$, $\gamma = 0.0153$, $\delta = 0.035$, $\theta = 0.32$, and $\psi = 0.58$. While the capital share appears to be on the high side and the depreciation rate on the low side as compared to international evidence, these values were calculated by the authors from available data. The depreciation rate is calculated using data on the capital stock estimates from the United States National Accounts from valuation of past investments for each type of unit of capital such that $\delta = 1 - \frac{K_{t+1} - I_t}{K_t}$. Capital's factor share was adjusted so that the share of capital reproduced the observed income of capital for Brazil in the late 1990s using the Brazilian National Income Accounts. (However, Bugarin et al. (2002) look at the Brazilian National Income Accounts and identify the capital share in the economy to be about 50 percent of output. The authors acknowledge this measure seems to be quite high, in particular considering the significant unreported income generated by self-employed and family workers, common features of Latin American countries. Rather the authors use information from a survey on household income and approximate the capital share in the Brazilian economy to 35 percent.) The rate of technological progress was obtained by adjusting an exponential trend to the United States output per worker series, correcting for the increase in the average schooling of the labor force.

Parameter values for the CES production function follow Ellery-Jr. (2012), including: $\mathrm{a} = 0.4$ and $\sigma = 1.67$. It is noteworthy that the value for the elasticity of substitution is assumed to be greater than 1, in contrast to the standard assumption that it is less than or equal to 1. It measures the extent to which firms can substitute capital for labor as the relative productivity or the relative cost of the two factors changes. When this number is large, it means that firms can easily substitute between capital and labor. If this number is above 1, then a given percentage change in will exceed the associated percentage change in. For example, an increase in the capital stock would raise the capital-labor ratio but lower the wage-capital return ratio by a smaller percentage, hence the share of capital in total income would rise as the capital-labor ratio increased. This value is taken from Murata and Lopes (2007) who calculate the elasticity of substitution in Brazil after the stabilization of the real.

## c. Results

The results of the newly constructed capital-output ratio, TFP (from both the Cobb-Douglas and CES production functions), and PTFD series are illustrated in Figure 7. Constructing a new capital series using Brazil's published nominal gross fixed capital formation and Brazil's published fixed capital formation price deflator results in a capital-output ratio in 2006 that is 34 percent above its level in 1989. In contrast, a declining capital-output ratio between 1989-2006 is observed when using the newly constructed capital price indexes to deflate Brazil's nominal gross fixed capital formation series for each of the definitions of capital. When deflating Brazil's nominal gross fixed capital formation using the capital price index according to the BEC definition of capital, the capital-output ratio is 38 percent lower in 2006 than in 1989. Using the capital price index according to the SITC7 and BEA definitions of capital, the capital-output ratio is 28 percent and 31 percent lower in 2006 than in 1989, respectively. These results also contrast sharply with documented large increases in the capital-output ratio for Brazil during this time period when using the published capital series by IPEA (see, for example, Gomes, Bugarin and Ellery-Jr (2005)).[37]

Brazil's newly constructed TFP and PTFD series also behave differently. Due to the years for which trade data exist, we acknowledge that we cannot explain the negative growth of TFP in the 1980s, as we still observe significant falls in Brazil's TFP between 1989 and 1992. When using a Cobb-Douglas production function, the level of TFP in 1992 is about 11 percent lower than in 1989 for each of the series constructed using the new capital price indexes. This fall is not as dramatic as that observed when using Brazil's fixed capital formation price series, being about 16 percent lower in 1992 than in 1989. However, this series shows no recovery in TFP between 1992 and 2006, resulting in an estimated TFP level about 15 percent lower in 2006 than in 1989. In contrast, our estimated TFP series constructed using the new capital

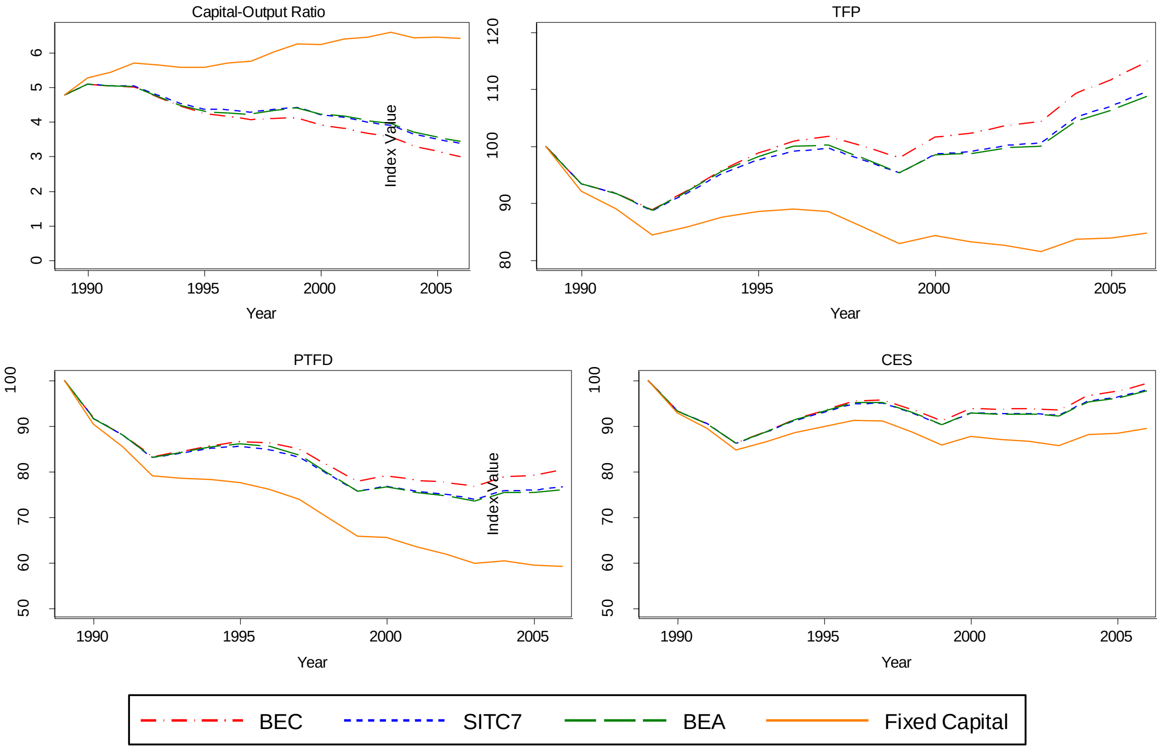
[37] Note that Gomes, Bugarin and Ellery-Jr (2005) calculate a fairly constant capital-output ratio for Brazil over the periods 1950-2000 by assuming that a fixed proportion of investment is wasted and not turned into capital by including waste in the perpetual inventory method. The authors do not provide TFP estimates using this new series of capital.

price indexes document a significant recovery in TFP. Between 1992 and 2006, the cumulative increase observed in Brazil's TFP constructed following the BEC definition of capital is 30 percent, following the SITC7 definition is 22 percent, and the BEA definition is 25 percent. This amounts to an average annual growth rate in TFP of 1.9 percent, 1.5 percent, and 1.6 percent, respectively. Overall, the level of TFP in 2006 is 15 percent higher than in 1989 when using the capital price index according to the BEC definition of capital, 8 percent using the SITC7definition, and 11 percent using the BEA definition.

When using a CES production function, the TFP series for each of the definitions of capital recovers between 1992 and 2006, however not to 1989 levels. Additionally, the recovery of the TFP stops very abruptly in 1996, as opposed to the recovery of the TFP series estimated using the Cobb-Douglas production function that continues through 2006. Eller-Jr. (2012) provides one potential explanation for this finding by separating capital and labor productivity in a CES production function. While overall TFP remains fairly constant, capital and labor productivity show very different trends. Post-1996, capital productivity continues to increase until 2003 while labor productivity declines quite dramatically. After 2003 the opposite result holds, and the recovering of labor productivity is driven largely by the accumulation of human capital throughout this period in Brazil. Allowing for substitution between capital and labor to reflect changes in their rates of returns could be one reason the TFP series remains flat.

After accounting for the evolution of the technological frontier, Brazil's discounted TFP (PTFD, or Brazil's country-specific productivity) still falls between 1989 and 2006 when using the newly constructed capital price indexes. Yet these series are fairly stable after 1992. Furthermore, this decrease is not as dramatic when comparing the estimates using Brazil's fixed capital formation price deflator, which continues to decline between 1992 and 2006. However, the inclusion of human capital is primarily responsible for the differences between the TFP series and the discounted TFP (PTFD) series, suggesting that much of the TFP improvement is being driven by human capital or labor productivity. Similar results are found in Eller-Jr. (2012). The results suggest that Brazil's TFP has progressed at a rate nearly similar to that of the technological frontier since 2002, which is different from what is found in Gomes, Pessôa and Veloso (2003). These results suggest that external links with other market economies through trade and investment are important drivers of productivity growth in Brazil.

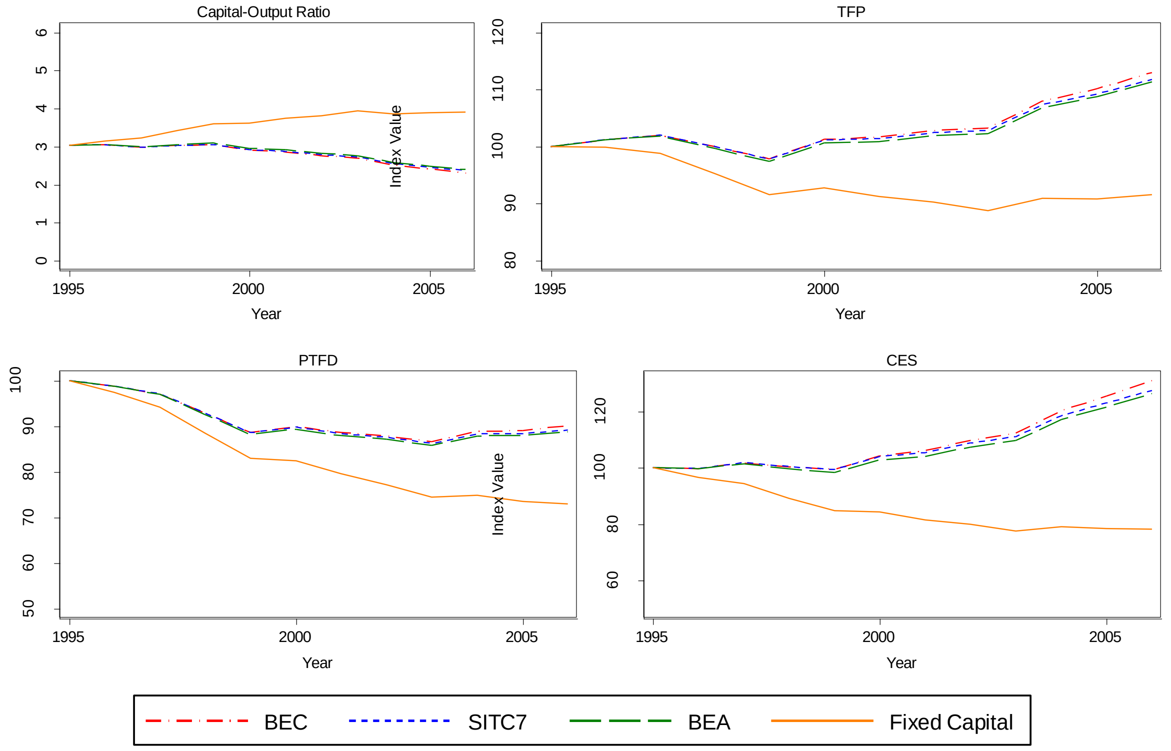
**Figure 7: Results, 1989-2006**




Notes: Figure 7 plots the TFP series constructed using each of the three definitions of capital goods (BEA in red, BEC in blue, SITC7 in green) and Brazil's published fixed capital price deflator (orange) to deflate the investment series prior to constructing the capital series using the perpetual inventory method from 1989-2006. The TFP is based to equal 100 in the base year (1989).

Furthermore, the results hold when restricting the sample to 1995-2006 (Figure 8), after stabilization of the real. When using either a Cobb-Douglas or CES production function, TFP in Brazil is higher than levels achieved in the mid-1990s. Thus Brazil's low productivity performance after the trade and economic reforms is in part an illusion due to mis-measurement. Brazil is shown to be more productive than the existing Brazilian literature has been arguing and Brazil is more technologically advanced today than it was 20 years ago.

**Figure 8: Results, 1995-2006**




Notes: Figure 8 plots the TFP series constructed using each of the three definitions of capital goods (BEA in red, BEC in blue, SITC7 in green) and Brazil's published fixed capital price deflator (orange) to deflate the investment series prior to constructing the capital series using the perpetual inventory method from 1995-2006. The TFP is based to equal 100 in the base year (1995).

## 7. Concluding Remarks

In this study we correct for mis-measurement of the price of capital by constructing a new capital price index using international trade data on investment goods' prices, then adjust the index to reflect domestic Brazilian prices using the real (R$)/US dollar ($) nominal exchange rate and Brazilian tariff data. The new price series behave completely different than the deflators published in the Brazilian national accounts. The percentage difference between the constructed capital price indexes and the published GDP deflator ranges from 23 to 128 percent and the published fixed capital formation price deflator from 48 to 173 percent.

With the capital price index in hand, a new capital stock series is computed and used to replicate simple estimates of TFP. We document a fall in Brazil's capital-output ratio, as opposed to significant capital deepening observed in the literature during this time period. Finally, our results show a significant recovery in Brazil's TFP between 1992 and 2006 when using the newly constructed capital price indexes to deflate the capital series, with a cumulative increase ranging between 22 and 30 percent. This amounts to an average annual growth rate in TFP ranging between 1.5 and 1.9 percent. Overall, the level of TFP in 2006 is between 8 and 15 percent higher than in 1989. The results show that it is the divergence in the relative price of capital to GDP that drives the results, not simply hyperinflation.

## Works Cited


Alfaro, Laura and Faisal Z. Ahmed (2007). "The Price of Capital: Evidence from Trade Data." Harvard Business School Working Paper WP 07-1073, Allston, MA.

Alfaro, Laura and Eliza Hammel (2007). "Capital Flows and Capital Goods." *Journal of International* Economics, 72(1): 128-150.

Banco Central do Brasil (2012). Time Series Management System. Retrieved from https://www3.bcb.gov.br/sgspub/localizarseries/localizarSeries.do?method=prepararTelaLocalizarSeries.

Bergstrand, Jeffrey H. (1983). "Measurement and Determinants of Intra-industry International Trade." In <u>Intra-Industry Trade: Empirical and Methodological Aspects</u>, edited by P.K.M. Thakaran, Elsevier, Amsterdam: 201-253.

Bergstrand, Jeffrey H. (1990). "The Heckscher-Ohlin-Samuelson Model, The Linder Hypothesis and the Determinants of Bilateral Intra-Industry Trade." *The Economic Journal*, 100(403): 1216-1229.

Barro, Robert and Jong-Wha Lee (2010), "A New Data Set of Educational Attainment in the World, 1950-2010." NBER Working Paper 15902, National Bureau of Economic Research, Cambridge, MA. Available at http://www.barrolee.com/data/dataexp.htm.

Bonelli, Regis and Renato Fonseca (1998). "Ganhos de Produtividade e de Eficiência: Novos Resultados Para a Economia Brasileira." *Pesquisa e Planejamento Econômico*, 28(2): 273-314.

Brazilian Institute of Geography and Statistics. Available at http://www.ibge.gov.br/english/.

Bugarin, Mirta N. S., Roberto de Goes Ellery-Jr, Victor Gomes, and Arilton Teixeira (2002). "The Brazilian Depression in the 1980s and 1990s." Mimeograph, Instituto de Pesquisa Economica Aplicada, Brasília. (As cited in Muendler, Servén and Sepúlveda (2004).) Revised version published in Great Depressions of the Twentieth Century, edited by Timothy Kehoe and Edward C. Prescott (2007), Minneapolis: Federal Reserve Bank of Minneapolis.

Center for International Data at UC Davis, U.S. Import and Export Data. Available at http://cid.econ.ucdavis.edu.

de Carvalho Filho, Irineu and Marcos Chamon (2012). "The myth of post-reform income stagnation: Evidence from Brazil and Mexico." *Journal of Development Economics*, 97: 368-86.

de Long, J. Bradford and Lawrence H. Summers (1991). "Equipment Investment and Economic Growth." *The Quarterly Journal of Economics*, 106(2): 445-502.

Eaton, Jonathan and Samuel Kortum (2001). "Trade in Capital Goods." *European Economic Review*, 45(7): 1195-1235.

Ellery-Jr., Roberto (2012). "Produtividade Total dos Fatores no Brasil no Período Pós-Reformas: Comportamento e Possiveis Explicações."

Feenstra, Robert C. (1996). "U.S. Imports, 1972-1994: Data and Concordances." NBER Working Paper No. 5515, National Bureau of Economic Research, Cambridge, MA.

Feenstra, Robert C., Robert E. Lipsey, and Harry P. Bowen (1997). "World Trade Flows, 1970-1992, with Production and Tariff Data." NBER Working Paper No. 5910, National Bureau of Economic Research, Cambridge, MA.

Feenstra, Robert C., John Romalis, and Peter K. Schott (2002). "U.S. Imports, Exports, and Tariff Data, 1989-2001." NBER Working Paper No. 9387, National Bureau of Economic Research, Cambridge, MA.

Ferreira, Pedro Cavalcanti, Roberto Ellery-Jr., and Victor Gomes (2008). "Produtividade Agregada Brasileira (1970-2000): Declínio Robusto e Fraca Recuperação." *Est. Econ.*, 38(1): 31-53.

Frensch, Richard and Vitalija Gaucaite Wittich (2009). "Product Variety and Technical Change." *Journal of Development Economics*, 88: 242-257.

Gomes, Victor, Mirta N. S. Bugarin, and Roberto Ellery-Jr. (2005). "Long-run Implications of the Brazilian Capital Stock and Income Estimates." *Brazilian Review of Econometrics*, 25(1): 67-88.

Gomes, Victor, Samuel de Abreu Pessôa, and Fernando A. Veloso (2003). "Evolução da Produtividade Total Dos Fatores Na Economia Brasileira: Uma Análise Comparativa." *Pesquisa e Planejamento Econômico*, 33(3): 389-434.

Hall, Robert E. and Charles I. Jones (1999). "Why Do Some Countries Produce So Much More Output Per Worker than Others?" *The Quarterly Journal of Economics*, 114(1): 83-116.

Harberger, Arnold C. (1978). "Perspectives of Capital and Technology in Less Developed Countries." In Contemporary Economic Analysis, edited by M.J. Artis and A.R. Nobay, London: Croom Helm.

Hiratsuka, Daisuke (2008). "Production Fragmentation and Networks in East Asia Characterized by Vertical Specialization." In Vertical Specialization and Economic Integration in East Asia, edited by Daisuke Hiratsuka and Yoko Uchida, Chosakenkyu-Hokokusho, IDE-JETRO.

International Monetary Fund (2004). Consumer Price Index Manual: Theory and Practice, Washington, DC.

International Monetary Fund (2009). Export and Import Price Index Manual: Theory and Practice, Washington, DC.

King, Robert G. and Ross Levine (1994). "Capital Fundamentalism, Economic Development, and Economic Growth." *Carnegie-Rochester Conference on Public Policy*, 40: 259-292.

Morandi, Lucilene (1998). "Estoque de riqueza e a poupanca do setor privado no Brasil." Texto para Discussão 572, Instituto de Pesquisa Econômica Aplicada, Rio de Janeiro, Brazil.

Morandi, Lucilene and Eustáquio J. Reis (2003). "Estimativa do Estoque de Capital Fixo-Brasil, 1950-2000." Instituto de Pesquisa Econômica Aplicada, Rio de Janeiro, Brazil.

Muendler, Marc, Luis Servén, and Claudia Sepúlveda (2004). "Productivity growth in Brazilian industry." In Brazil: The New Growth Agenda, Volume II, Detailed Report, edited by Mark R. Thomas and Indermitt Gill, Washington, DC: World Bank.

Murato, M. and R. Lopes (2007). Elasticidade de Substituição Entre Capital e Trabalho na Economic Brasileria: Uma análise de dados em painel. *A Economia em Revista*, 15(2).

Nehru, Vikram and Ashok Dhareshwar (1993). "A New Database on Physical Capital Stock: Sources, Methodology and Results." *Revista de Analisis Economico*, 8(1): 37-59.

Organization for Economic Co-Operation and Development (2008). Handbook on Constructing Composite Indicators: Methodology and User Guide, Paris.

Pritchett, Lant (2000). "The Tyranny of Concepts: CUDIE (Cumulated, Depreciated, Investment Effort) Is *Not* Capital." *Journal of Economic Growth*, 5: 361-384.

Pinheiro, Armando Castelar, Indermit S. Gill, Luis Servén, and Mark Roland Thomas (2001). "Brazilian Economic Growth, 1900-2000: Lessons and Policy Implications." IDB Publications No. 14838, Inter-American Development Bank, Washington, DC.

Receita Federal (2011). Available from http://www.receita.fazenda.gov.br/aliquotas/DownloadArqTIPI.htm.

Sakellaris, Plutarchos and Focco W. Vijselaar (2004). "Capital Quality Improvement and the Sources of Growth in the Euro Area." Working Paper Series No. 368, European Central Bank, Frankfurt.

Silva Filho, TN Teixeira da (2001). "Estimando o Produto Potencial Brasileiro: Uma Abordagem de Função de Produção." Working Paper No. 17, Banco Central do Brasil, Brazília.

Turkcan, Kemal (2007). "Outward Foreign Direct Investment and Intermediate Goods Exports: Evidence from the USA." *Economie Internationale*, 4: 51-71.

United Nations Statistics Division, United Nations Classification by Broad Economic Categories. Retrieved from http://unstats.un.org/unsd/cr/registry/regcst.asp?Cl=10&Lg=1.

United Nations Statistics Division, United Nations Classifications Publications. Retrieved from http://unstats.un.org/unsd/cr/registry/regdnld.asp?Lg=1.

World Bank (2011). World Development Indicators. Retrieved from http://data.worldbank.org/indicator.

World Trade Organization (2011). Regional Trade Agreements Database. Retrieved from http://www.wto.org/english/tratop_e/region_e/region_e.htm.

Xu, Bin and Jianmao Wang (1999). "Capital Goods Trade and R&D Spillovers in the OECD." *The Canadian Journal of Economics*, 32(5): 1258-1274.

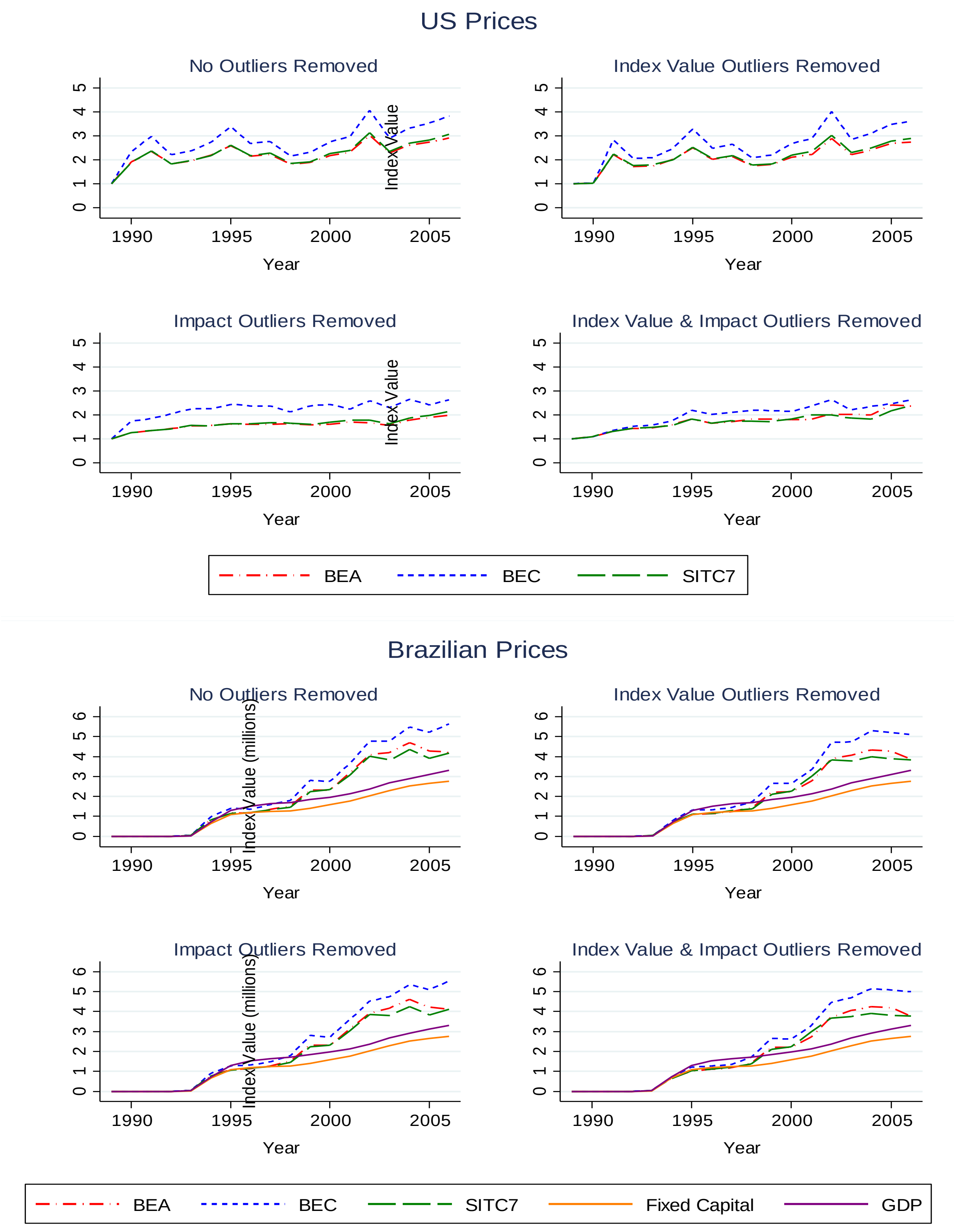
Xu, Bin and Jianmao Wang (2000). "Trade, FDI, and International Technology Diffusion." *Journal of Economic Integration*, 15(4): 585-601.

# Appendix

## a. Results of the Capital Price Index

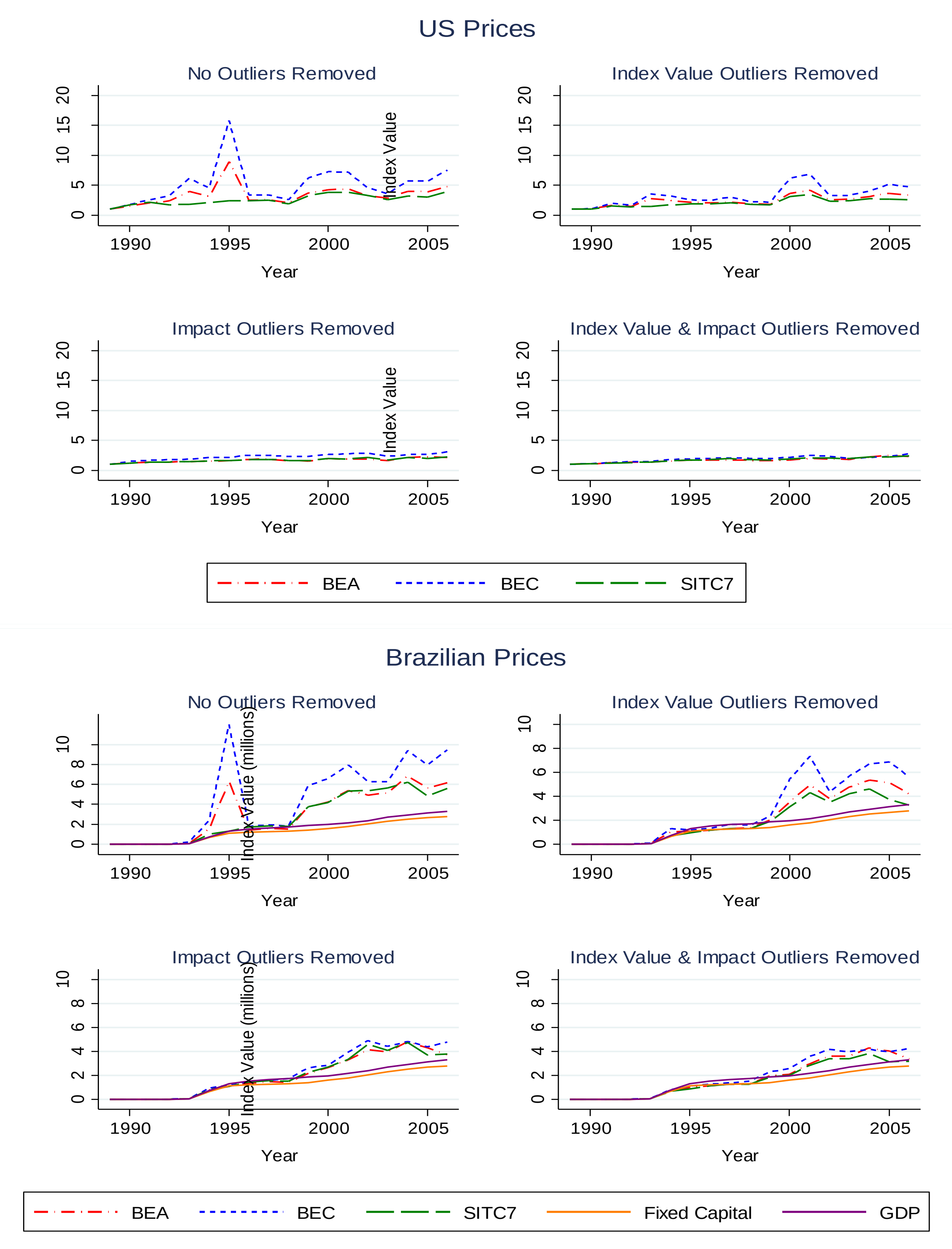
Presented below is the constructed capital price index for each of the definitions of capital goods for each index methodology. For each index methodology, the results are presented in nominal United States prices as well as nominal Brazilian prices. The indexes are also plotted against Brazil's fixed capital formation price deflator (Brazilian Institute of Geography and Statistics 2012) and Brazil's GDP deflator (Brazilian Institute of Geography and Statistics 2012), where appropriate. Furthermore, each index is presented without any outliers removed (top left), statistical checking of outliers removed (top right), impact outliers removed with a 10 percent impact threshold (bottom left), and both statistical checking and impact outliers removed (bottom right).

**Figure A.1: Capital Price Index, 1989-2006, Base Index**



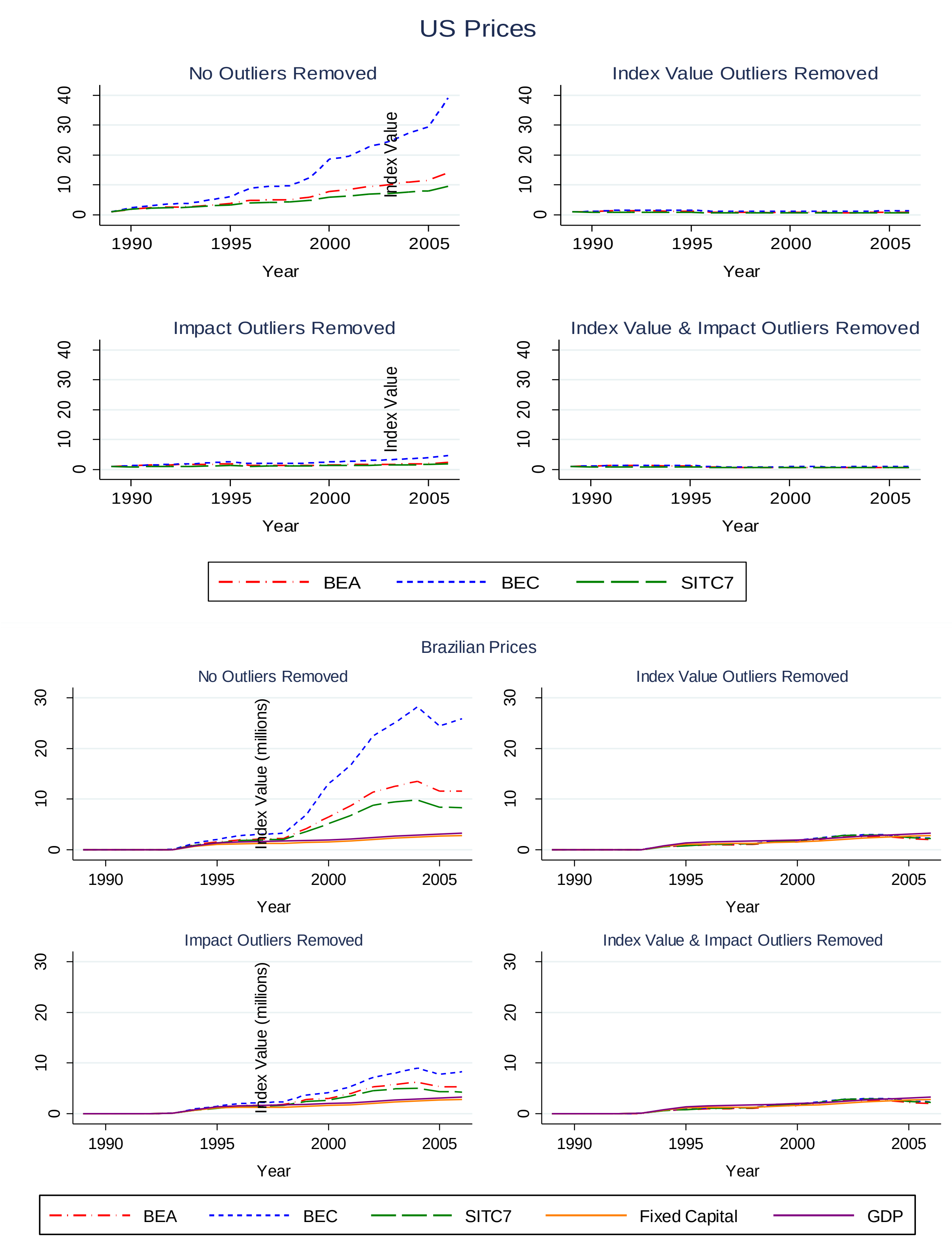

Notes: Figure A.1 plots the capital price index using base index methodology in US prices (top panel) and Brazilian prices (bottom panel) for the three definitions of capital goods (BEA in red, BEC in blue, SITC7 in green) and Brazil's published price deflators (fixed capital in orange, GDP in purple) from 1989-2006 using different methodologies to remove outliers. The index is based to equal 1 in the base year (1989).

**Figure A.2: Capital Price Index, 1989-2006, Spliced Index**




Notes: Figure A.2 plots the capital price index using spliced index methodology in US prices (top panel) and Brazilian prices (bottom panel) for the three definitions of capital goods (BEA in red, BEC in blue, SITC7 in green) and Brazil's published price deflators (fixed capital in orange, GDP in purple) from 1989-2006 using different methodologies to remove outliers. The index is based to equal 1 in the base year (1989).

**Figure A.3: Capital Price Index, 1989-2006, Chain Index**

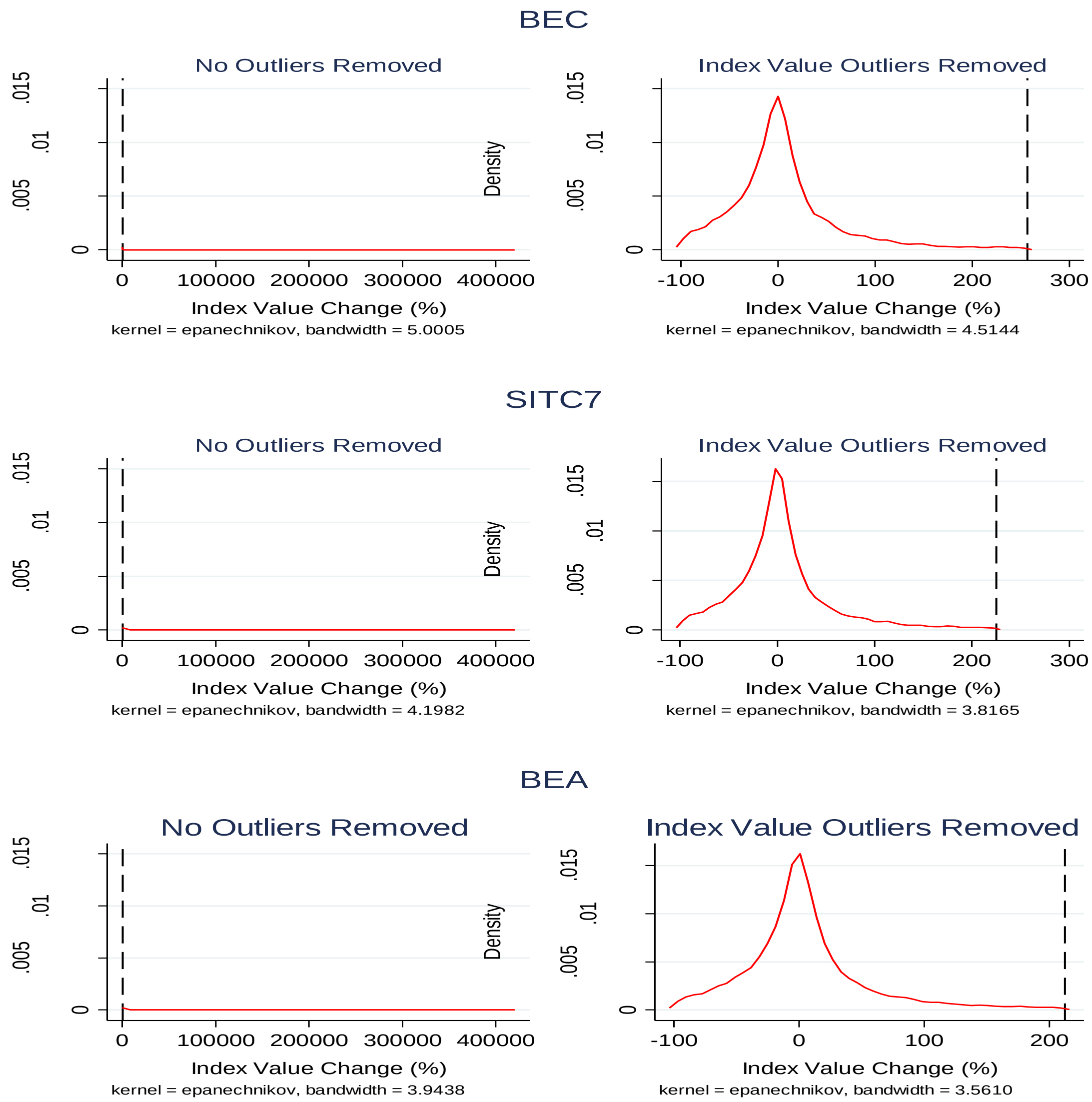



Notes: Figure A.3 plots the capital price index using chain index methodology in US prices (top panel) and Brazilian prices (bottom panel) for the three definitions of capital goods (BEA in red, BEC in blue, SITC7 in green) and Brazil's published price deflators (fixed capital in orange, GDP in purple) from 1989-2006 using different methodologies to remove outliers. The index is based to equal 1 in the base year (1989).

## b. Kernel Density Plots

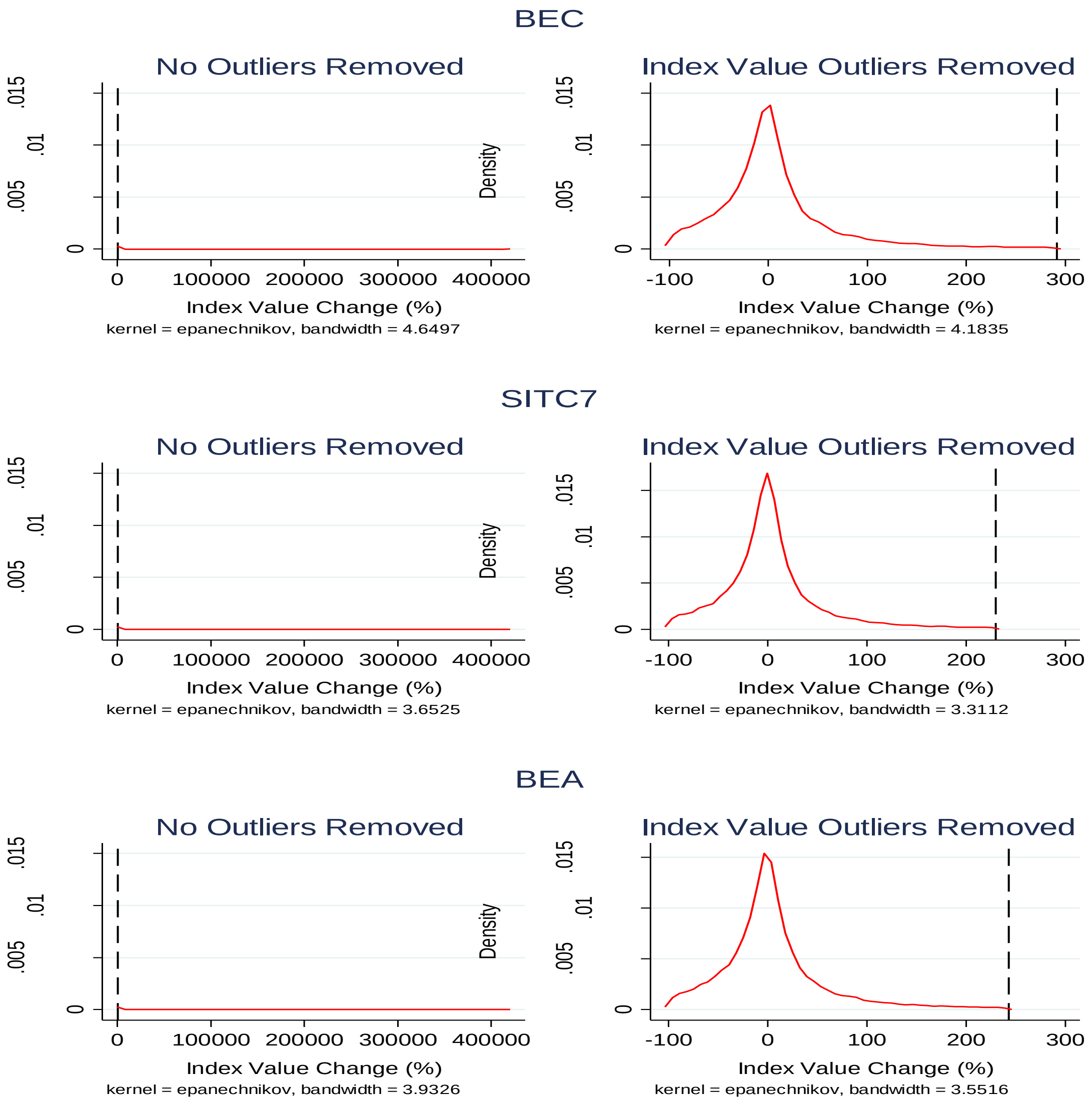
Kernel density plots are provided to approximate the distribution of the percentage change in unit values to identify outliers under the statistical checking of input data approach. The distributions are clearly significantly left skewed, and as such, the top 2.5 percent of the distribution were identified as outliers and dropped from the sample. A plot is provided for before the outliers are removed (left) as well as after the outliers are removed (right).

**Figure B.1: Kernel Density Plots, Base Index**



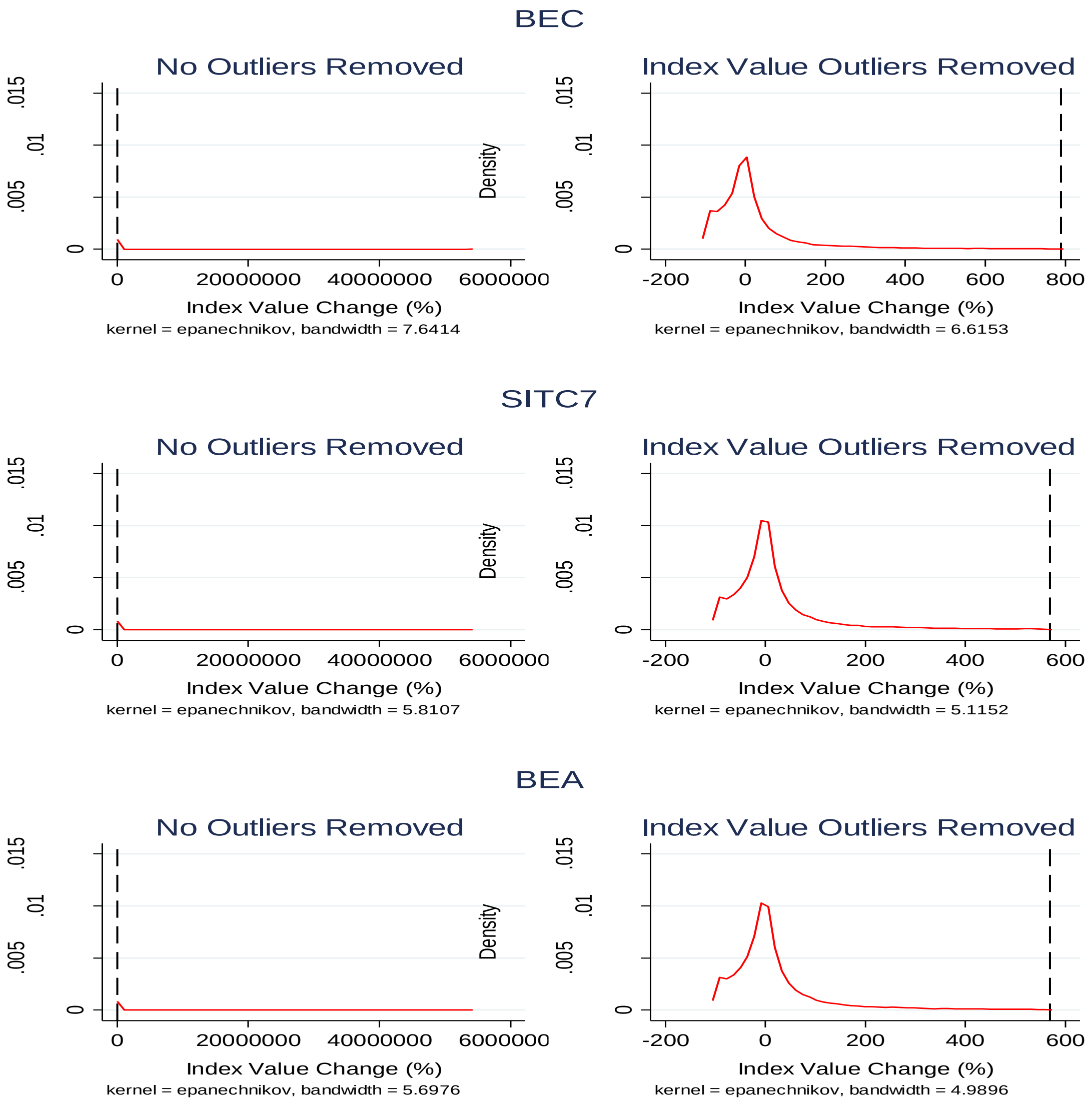

Notes: Figure B.1 plots the probability distribution of unit values with no outliers removed (left panel) and with the top 5 percent of index values removed (right panel) using base index methodology for the three definitions of capital goods (BEA, BEC, SITC7).

**Figure B.2: Kernel Density Plots, Spliced Index**




Notes: Figure B.2 plots the probability distribution of unit values with no outliers removed (left panel) and with the top 5 percent of index values removed (right panel) using spliced index methodology for the three definitions of capital goods (BEA, BEC, SITC7).

**Figure B.3: Kernel Density Plots, Chain Index**

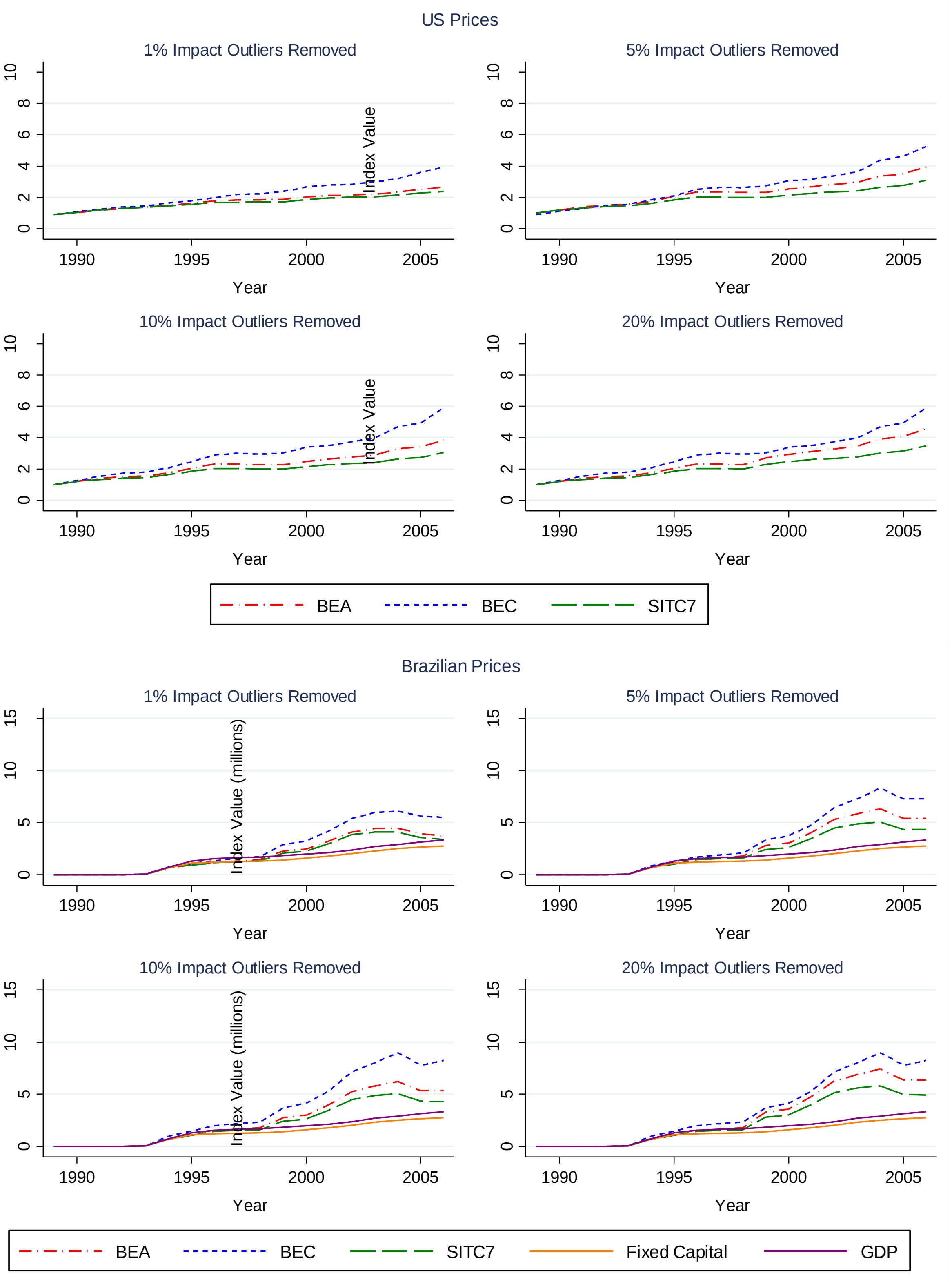



Notes: Figure B.3 plots the probability distribution of unit values with no outliers removed (left panel) and with the top 5 percent of index values removed (right panel) using chain index methodology for the three definitions of capital goods (BEA, BEC, SITC7).

### c. Impact Outlier Thresholds

This section presents the chain index expressed in United States prices and Brazilian prices with impact outliers removed at different impact thresholds, including 1 percent, 5 percent, 10 percent, and 20 percent.

**Figure C.1: Chain Index with 1%, 5%, 10% and 20% Impact Thresholds**



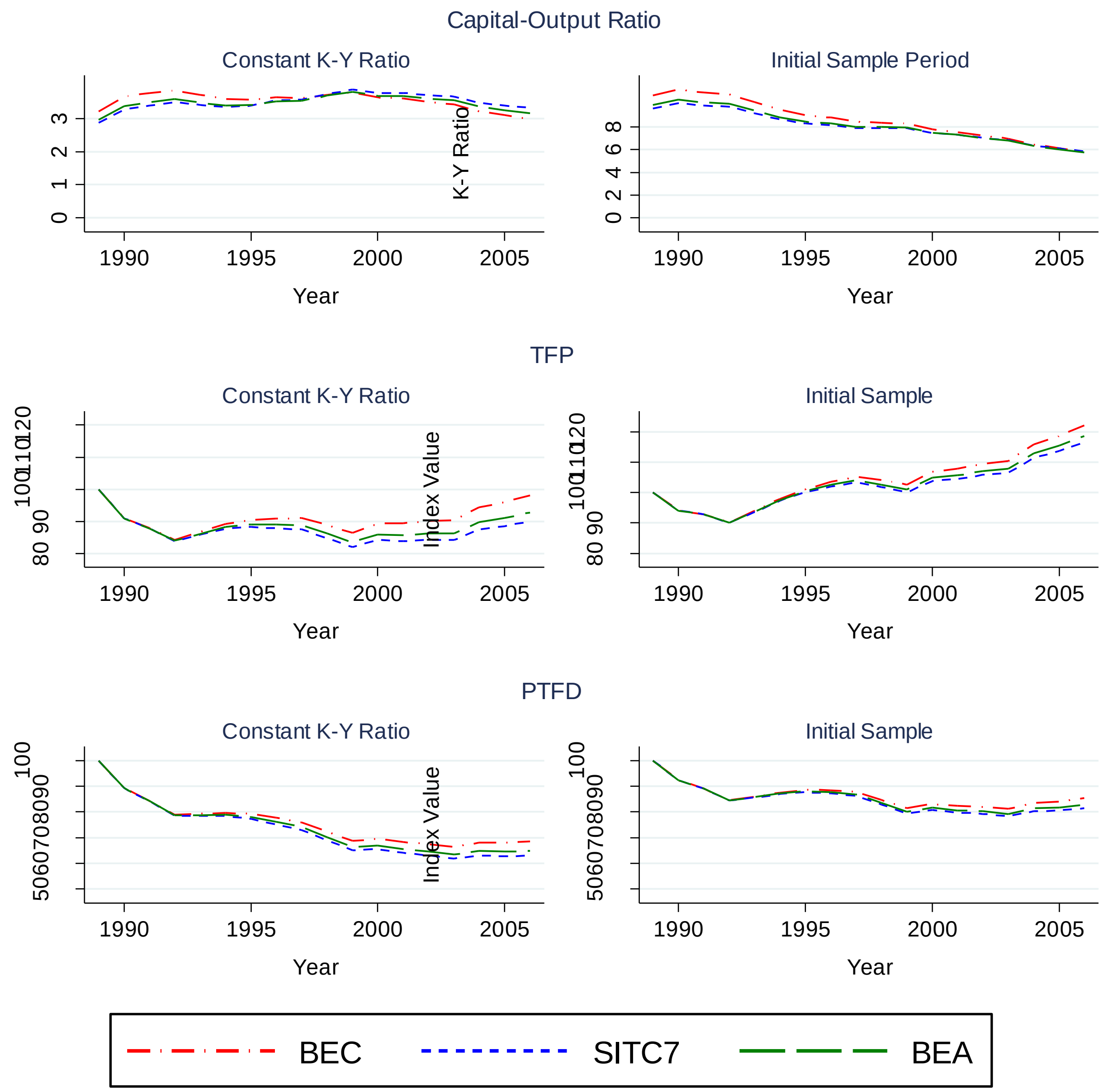

Notes: Figure C.1 plots the capital price index using chain index methodology in US prices (top panel) and Brazilian prices (bottom panel) for the three definitions of capital goods (BEA in red, BEC in blue, SITC7 in green) and Brazil's published price deflators (fixed capital in orange, GDP in purple) from 1989-2006 using an impact threshold approach to identify outliers for different thresholds (1%, 5%, 10%, 20%). The index is based to equal 1 in the base year (1989).

## d. Alternative Estimates of the Initial Level of Capital

### i. Growth rate of output

As suggested above, the initial capital stock can be estimated using the growth rate of output over the initial sample period 1989-1993 as well as the time period with a constant capital output ratio 1967-1976. The capital-output ratio, TFP, and PTFD series were then calculated using these capital series, presented in Figure D.1.

**Figure D.1: Constant K-Y Period and Initial Sample Period**




Notes: Figure D.1 plots the capital output ratio (top panel), TFP (middle panel), and detrended TFP (bottom panel) series constructed using each of the three definitions of capital goods (BEA in red, BEC in blue, SITC7 in green) to deflate the investment series prior to constructing the capital series using the perpetual inventory method from 1989-2006. The TFP is based to equal 100 in the base year (1989). In the left panels the initial capital stock is estimated using the growth rate of output over the period with a constant capital output ratio

1967-1976. In the right panels the initial capital stock is estimated using the growth rate of output over the initial sample period 1989-1993.

### ii. Growth rate of population and technological progress

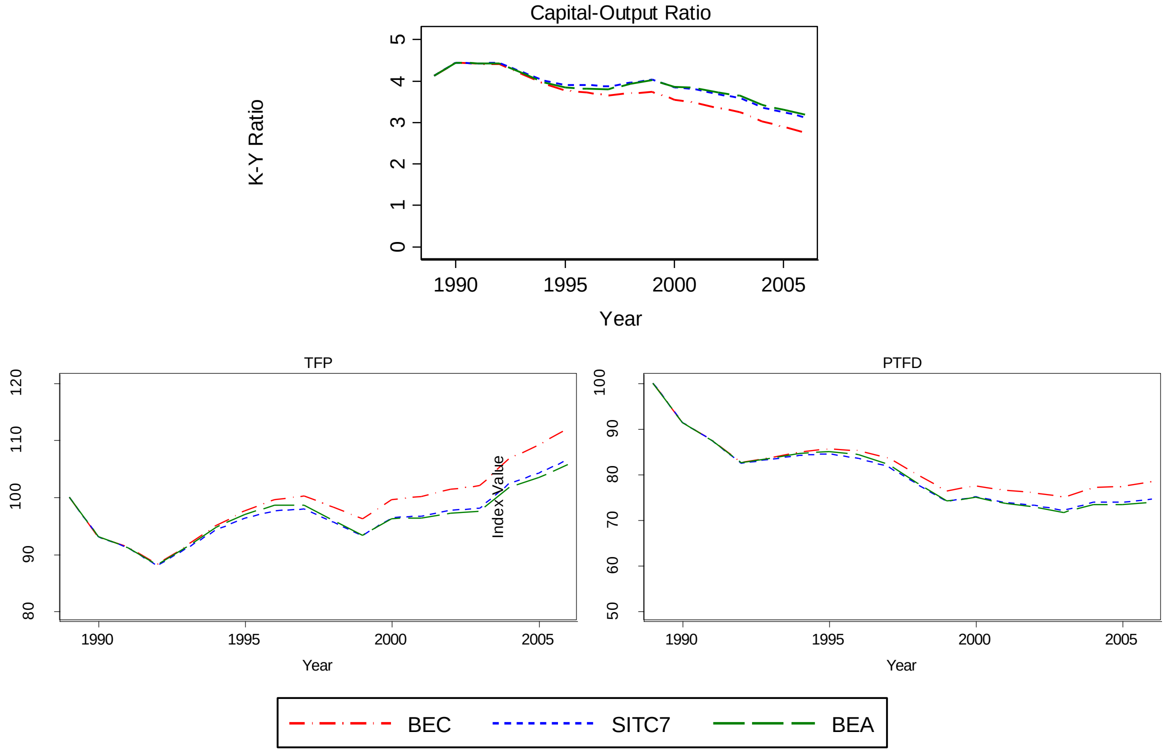
Alternatively, since in the steady state capital and thus output grow at the rate of population growth plus the growth rate of technological progress, the equation for the initial capital stock can be rewritten as

$$K_0 = \frac{I_1}{(1+n)(1+\gamma)-(1-\delta)}$$

where $n$ is the rate of growth rate of the population and $\gamma$ is the rate of technological progress. Figure D.2 presents the corresponding capital-output ratio, TFP, and PTFD series using this initial level of the capital stock. This approach was used by Gomes, Pessôa and Veloso (2003).

Average population growth was downloaded from the World Bank's World Development Indicators: Population growth (annual percent), defined as "Annual population g rate for year t is the exponential rate of g of midyear population from year t-1 to t, expressed as a percentage. Population is based on the de facto definition of population, which counts all residents regardless of legal status or citizenship--except for refugees not permanently settled in the country of asylum, who are generally considered part of the population of the country of origin." This was then averaged over the sample period to obtain the average population growth rate.

**Figure D.2: Growth Rate of Population and Technological Progress**



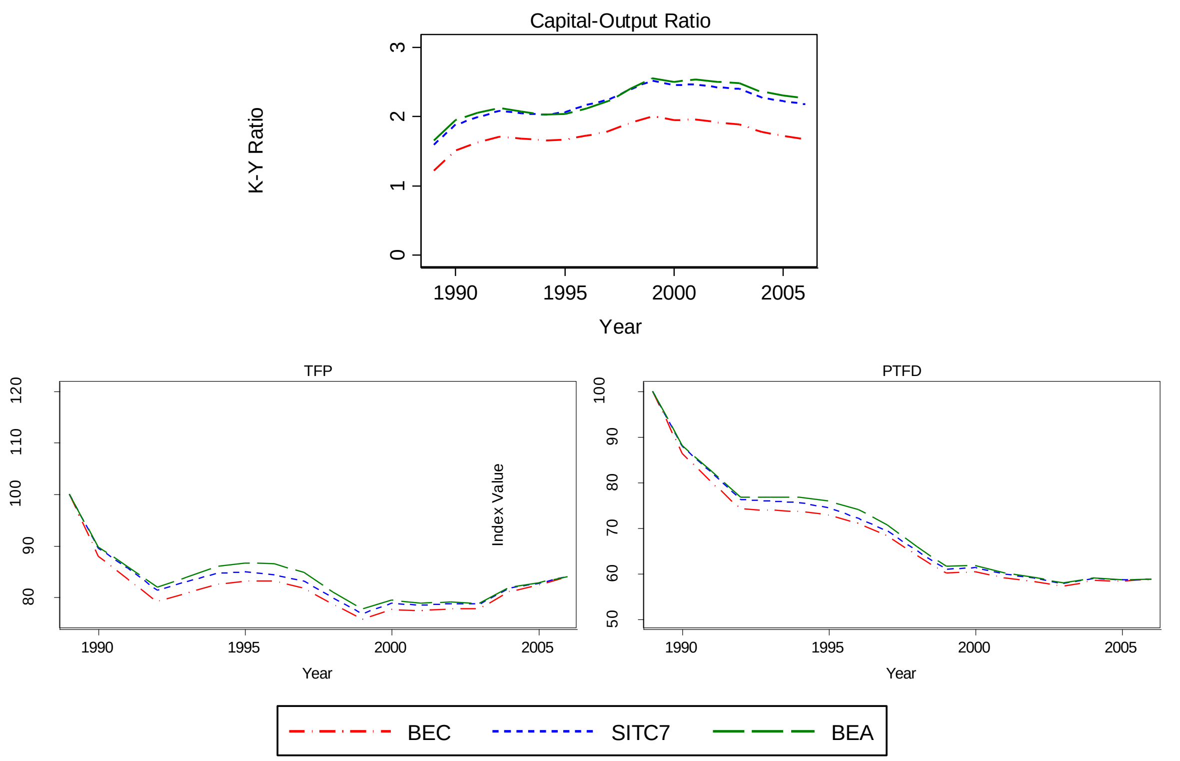

Notes: Figure D.2 plots the capital output ratio, TFP, and detrended TFP series constructed using each of the three definitions of capital goods (BEA in red, BEC in blue, SITC7 in green) to deflate the investment series prior to constructing the capital series using the perpetual inventory method from 1989-2006. The TFP is based to equal 100 in the base year (1989). The initial capital stock is estimated using the growth rate of population and technological progress.

### iii. Current capital-output ratio

A quick approach as suggested by Nehru and Dhareshwar (1993) is to assume that the capital-output ratio in the initial year is the same as the capital-output ratio in the last year of the investment series data. The capital-output ratio in the last year can be calculated using the capital stock generated by the PIM assuming an initial capital stock of zero.[38] The PIM is then re-applied using this initial level of capital. Figure D.3 presents the corresponding capital-output ratio, TFP, and PTFD series.

[38] Or, the average capital-output ratio in the last few years may provide a better estimate.

**Figure D.3: Final Capital-Output Ratio**




Notes: Figure D.3 plots the capital output ratio, TFP, and detrended TFP series constructed using each of the three definitions of capital goods (BEA in red, BEC in blue, SITC7 in green) to deflate the investment series prior to constructing the capital series using the perpetual inventory method from 1989-2006. The TFP is based to equal 100 in the base year (1989). The initial capital stock is estimated using the current capital-output ratio.

# Chapter 4

# Monitoring Export Vulnerability to Changes in Growth Rates of Major Global Markets

Claire H. Hollweg
with Daniel Lederman
and José-Daniel Reyes

Hollweg, C.H., Lederman, D. & Reyes, J-D. (2012) Monitoring export vulnerability to changes in growth rates of major global markets.
*The World Bank, Policy Research Working Paper No. 6266, November, pp. 1-26*