Universidad de la República
Departamento de Economía - Facultad de Ciencias Sociales
Doctorado en Economía

Three Essays about Migration, Gender and Family Economics

Florencia Amábile Bruchou

Montevideo - Uruguay
February, 2022

Ciencias Sociales
Universidad de la República
URUGUAY

dECON
Facultad de Ciencias Sociales

Three Essays about Migration, Gender and Family Economics

Florencia Amábile Bruchou

Tesis de Doctorado presentada en el Doctorado en Economía - Departamento de Economía - Facultad de Ciencias Sociales de la Universidad de la República, como parte de los requisitos necesarios para la obtención del título de Doctor.

Director:
Ph.D. Prof. Rómulo A. Chumacero

Director Académico:
Ph.D. Prof. Álvaro Forteza

Montevideo - Uruguay
Febrero de 2022

Agradecimientos

Quisiera expresar mi más sincero agradecimiento a mi Director de Tesis Prof. Rómulo A. Chumacero, sin el cual esta tesis no hubiera sido posible. Su orientación, apoyo y crítica constructiva durante todo el proceso de doctorado resultan invaluables.

A Rafael le agradezco su amor, apoyo, paciencia y comprensión durante todos estos años. A Manuel por darnos el amor que nos faltaba y equilibrar nuestra vida. A mis padres, por su amor, por siempre alentarme en mis proyectos y por enseñarme a no rendirme hasta verlos cumplidos. A mi hermana, sobrinos, suegra y cuñados por todo su apoyo y cariño.

Agradezco el tiempo dedicado de compañeros de trabajo, colegas y amigos por el constante ánimo, las charlas, comentarios y sugerencias en el proceso de elaboración de esta tesis: Verónica Amarante, Ana Balsa, Marisa Bucheli, Maira Colacce, Rosario Domingo, Andrea Doneschi, Zuleika Ferre, Álvaro Forteza, Mariana Gerstenblüth, Cecilia González, Cecilia Noboa, Cecilia Olivieri, Ianina Rossi, Máximo Rossi, Gonzalo Salas, Graciela Sanroman, Victoria Tenenbaum, Patricia Triunfo, Andrea Vigorito y Leandro Zipitría.

A mis amigos de todas las horas, gracias por ser el soporte en este camino: Ignacio Álvarez, Analía Ardente, Natalia da Silva, Nora Gesto, Alejandra Marroig y Cecilia Oreiro.

Finalmente, quisiera reconocer el apoyo que recibí desde el Departamento de Economía de la Facultad de Ciencias Sociales de la Udelar, en especial a través de sus directores, quienes en todo momento me alentaron a seguir.

Gran parte del trabajo contenido en esta tesis ha sido financiado por la Beca de apoyo a docentes para estudios de posgrado en la Udelar de la Comisión Académica de Doctorado, durante el período 2017-2020.

A Rafael y Manuel

Resumen

Esta tesis consta de tres ensayos sobre migración, género y economía de la familia, en el marco de la economía aplicada y con especial énfasis en un país en desarrollo. El primer capítulo de mi tesis contribuye a la literatura de políticas fiscales y transferencias intergeneracionales. El segundo capítulo contribuye a la literatura sobre los efectos de la composición sexual de los hijos en el comportamiento de los padres y cómo puede afectar el desarrollo cognitivo y no cognitivo de los niños y niñas. El tercer capítulo contribuye a la literatura sobre género y economía de la familia.

El primer capítulo, en coautoría con Rómulo A. Chumacero, desarrolla un modelo de ciclos económicos de una economía pequeña y abierta a la movilidad de internacional del trabajo y agentes heterogéneos, con especial atención a los impuestos y las políticas de transferencia. La migración se produce como resultado del problema de maximización de las familias y, combinada con las remesas, hace posible la suavización del consumo. Este trabajo muestra cómo las transferencias del gobierno a los jóvenes y a los mayores, financiadas con impuestos distorsionantes, impulsan la migración de las personas en edad de trabajar y, entre éstas, a algunos de los miembros más calificados de la economía. El modelo se calibra para que coincida con la movilidad laboral en varios grupos de edad y nivel de calificación, así como con la dinámica del ciclo agregado de la economía uruguaya, incluidas las transferencias del gobierno y la volatilidad de la migración.

El segundo capítulo investiga cómo el sexo del segundo hermano afecta al desarrollo cognitivo y no cognitivo del primer hijo y al comportamiento de los padres cuando los niños tienen una media de 52 meses (y no más de 66 meses) en Uruguay. Dado que este trabajo analiza a los niños en edades muy tempranas, es poco probable que las influencias entre hermanos desempeñen un rol importante en los resultados cognitivos y no cognitivos, brindando una oportunidad para aislar los efectos que surgen puramente a través del comportamiento de los padres. La estrategia de identificación supera el sesgo en las preferencias de los padres sobre la composición del sexo de los hijos debido a la aleatoriedad del sexo

del segundo hijo. Los resultados muestran que tener un segundo hermano varón, relativo a tener una segunda hermana, afecta negativamente el desarrollo motor y el desarrollo no cognitivo del primer hijo, pero no afecta de forma diferenciada el desarrollo de la primer hija. Los principales mecanismos que podrían explicar las diferencias observadas entre los primeros hijos e hijas son: la menor probabilidad de vivir con ambos padres biológicos, la menor inversión de los padres en tiempo de calidad y la menor probabilidad de que asistan a la escuela pre-escolar.

El tercer capítulo analiza a través de la oferta de trabajo de hombres y mujeres, el poder de negociación al interior de hogares en parejas heterosexuales, que difieren en normas sociales respecto a la división del trabajo doméstico y que son clasificados como tradicionales, igualitarios y no tradicionales. Los datos de Uruguay muestran que los hogares son sensibles a cambios en el poder de negociación, medidos por la diferencia de ingresos no laborales entre los miembros de la pareja y el hecho de estar casados. Los resultados sugieren que un hombre relativamente más rico tiene mayor poder de negociación y ofrece menos horas al mercado de trabajo, y lo contrario ocurre con su pareja. Además, el hecho de estar casado reduce la oferta de trabajo de las mujeres y aumenta la de los hombres en los hogares igualitarios. Estos resultados son robustos a la corrección por sesgo de selección en el mercado laboral. Por último, los hogares menos normativos en materia de género asignan una mayor parte de los ingresos no laborales a las mujeres tras el proceso de negociación.

Capítulo 1:
Palabras clave: Uruguay, Migración, Agentes Heterogéneos, DSGE, Transferencias
Clasificación JEL: E27, F22, H31, I25

Capítulo 2:
Palabras clave: Uruguay, Hermanos, Composiciń sexual, Desarrollo infantil, Estructura familiar
Clasificación JEL: D19, J12, J13, J16

Capítulo 3:
Palabras clave: Uruguay, Modelos Collectivos de Oferta de Trabajo, Negociación intrahogar, Género, Normas Sociales
Clasificación JEL: D13, D91, J16, J22

Abstract

This thesis consists of three essays about migration, gender, and family economics, within the framework of applied economics and with a special focus on a developing country. The first essay of my thesis contributes to the literature on fiscal policies and intergenerational transfers. The second essay contributes to the literature on the effects of children's sex composition on parental behavior and how it may affect children's cognitive and non-cognitive development. The third essay contributes to the literature on gender and family economics.

The first chapter, co-authored with Rómulo A. Chumacero, develops a business cycle model of a small open economy with heterogeneous agents and international labor mobility, with a particular focus on taxes and transfer policies. Migration occurs as a result of the maximization problem of families and, combined with remittances, makes consumption smoothing possible. This paper shows how transfers from the government to young people and elders, funded with distortionary taxes, prompt the migration of people of working age and, among them, some of the most skilled members of the economy. The model is calibrated to match labor mobility in various age-skill groups and aggregate cycle dynamics of the Uruguayan economy, including government transfers and migration volatility.

The second chapter investigates how the gender of the second-born sibling affects first-born cognitive and non-cognitive development and their parental treatment when children are on average 52 months (and not older than 66 months) in Uruguay. Since the study looks at children at a very early age, sibling-to-sibling influences are unlikely to play a role and hence an opportunity to isolate effects that arise purely via parental treatment. The identification strategy overcomes parental preferences' bias of children's sex composition due to the randomness of the sex of the second-born child. Results show that first-born boys who have a same-sex younger sibling have lower levels of motor skills and non-cognitive development. In contrast, first-born girls are not affected by having a younger sister or brother. The main drivers of the differences between first-born boys and girls are the lower probability of boys living with both biological parents, less investment of parents in their

quality time, and a reduced likelihood of attending preschool.

The third chapter analyzes the impact of intra-household bargaining on the labor supply of heterosexual couples with different divisions of domestic work. The objective is to compare the decision-making process in families with egalitarian, traditional, and non-traditional gender role attitudes towards the division of domestic work. Data from Uruguay shows that couples of all types are sensitive to bargaining power shifts, as measured by the non-labor income difference between cohabiting partners and married couples. Results suggest that a relatively rich male has more bargaining power and supplies less labor, and the opposite is true for his partner. In addition, being married reduces the labor supply of women and increases that of men in egalitarian households. These results are robust to selection into employment correction. Finally, less gender-normative households assign a larger share of non-labor income to women after the negotiation process.

Chapter 1:
Keywords: Uruguay, Migration, Heterogeneous Agents, DSGE, Transfers
JEL Classification: E27, F22, H31, I25
Chapter 2:
Keywords: Uruguay, Siblings, Sex Composition, Childhood Development, Family Structure
JEL Classification: D19, J12, J13, J16
Chapter 3:
Keywords: Uruguay, Collective Models of Labor Supply, Intra-Household Bargaining, Gender, Social Norms
JEL Classification: D13, D91, J16, J22

Foreword

The three essays that compose this thesis are separate articles submitted or in preparation for submission to peer-reviewed scientific journals. I am the principal author of the three chapters presented in this thesis. The first chapter is the only one co-authored and was written jointly with my thesis supervisor Rómulo A. Chumacreo, Associate Professor in the Department of Economics at the University of Chile. This paper, entitled "Should I Stay or Should I Go?: The Economic Incentives of Intergenerational Taxes and Transfers in Uruguay" was accepted in December 2021 in the journal International Tax and Public Finance. The second chapter, entitled "Middle-term Effects of Sibling's Sex Composition on Early Childhood Development", has been submitted and rejected to the Journal of Population Economics. From this rejection, I have comments from two anonymous referees to introduce in a future version of the paper. The third chapter, entitled "Collective Labor Supply, Divisions of Domestic Work and Intra-Household Bargaining", is under preparation to be submitted to a peer-review journal.

Contents

List of Figures

List of Tables

Chapter 1

Should I Stay or Should I Go?: The Economic Incentives of Intergenerational Taxes and Transfers in Uruguay

1.1 Introduction

The effects of transfer programs on labor supply, household structure, intergenerational dependence, and migration have been a central issue in economic research (Moffitt, 1992). While these programs generally intend to improve the living conditions of the beneficiaries, they can have unintended consequences. This is especially true in economies with large migration flows, where the welfare state redistributes resources from the rich to the poor, and from middle-aged individuals to young people and the elderly (typically in cash or in-kind transfers). Whereas there is a vast literature on the linkages between fiscal policy and migration in developed (host) countries, there is a fewer discussion of its impact on developing (source) countries.[1]

The "brain drain" theory of migration suggests that high-skilled migration from developing countries erodes economic development in the source country due to losses in public investment (Berry and Soligo, 1969). Opposite to this view is the "brain gain" theory, which claims that labor mobility generates an "educational effect" where individuals with higher

[1]The global stock of migrants has grown from 153 million to 258 million people between 1990 and 2017; representing 3.4% of total population (United Nations, 2017).

expected income abroad accumulate additional human capital in the source country, counterbalancing the original "brain drain" (Stark and Taylor, 1991; Mountford, 1997). In turn, human capital networks, return migration, e.g., Stark et al. (1997), and remittances, e.g., Djajić (1986); Quibria (1997) can compensate (at least partly) for the income losses caused by the "brain drain". Quibria (1997) showed that the total effect of migration varies according to the volume of remittances, the distribution of factor endowments and the type of emigration.

This paper analyzes the incentives created through different tax and transfer policies in a small open economy, using a Dynamic Stochastic General Equilibrium (DSGE) model with international labor mobility. In this model, migration occurs as a result of the family's maximization problem, considering intertemporal consumption smoothing and the trade-off between migration costs and welfare gains from higher wages in the foreign economy.[2]

This work contributes to the literature in three main ways. Firstly, this is the first paper to use a DSGE model to evaluate fiscal policies and intergenerational transfers, with endogenous and heterogeneous labor migration (in terms of age and ability). This framework is more suitable than overlapping generations (OLG) models when analyzing the interaction between migration decisions, fiscal policies, and the economic structure. Second, this paper highlights the consequences of fiscal policies for labor supply and migration in developing countries. Third, although there is a vast literature on the "welfare magnet hypothesis" (Borjas, 1999), this work attempts to fill an existing gap in the literature, which is the role of the welfare state in developing countries when faced with huge levels of migration.[3]

This paper contrasts with the literature on migration due to the introduction of heterogeneous agents in age and ability on a DSGE model and the use of this framework to assess the interactions between intergenerational tax and transfer policies and migration. Some papers consider these interactions in an OLG framework, but almost none of them look at the effects on the source country, with Leers et al. (2004) being one of the exceptions. They analyze the incentives to migrate from an aging economy, where a Pay-As-You-Go (PAYG) social-security system is becoming unfunded. Using an OLG model, they conclude that

[2] Hanson (2010) shows that people who migrate to OECD countries from developing countries have higher levels of education, as compared to the people that do not. Similarly, Chiquiar and Hanson (2005) find that migrants from Mexico to the US are concentrated in the middle of the distribution in terms of education and wages.

[3] The "welfare magnet hypothesis" suggests that immigrants are more likely to move to countries where welfare systems are generous. Several studies have analyzed this hypothesis, but results are not conclusive (Giulietti, 2014). Jakubiak (2017) provides an extensive literature review on immigration and welfare systems, noticing that there is no discussion about the relationship between international migration and the redistributive policies in the countries of origin, which is the main focus of this paper.

when labor is heterogeneous in age and public pensions are affected by lobby from the elderly, there are additional incentives to emigrate to avoid the excessive taxes that fund the PAYG system. Other works that use an OLG framework focus on the effects of immigration and different migration policies in the host country on the sustainability of social security systems and their implications for fiscal policies (Sand and Razin, 2006, 2007; Storesletten, 2000).

Much of the literature that uses DSGE models to analyze migration is motivated by the aftermath of the Great Recession in Europe which, encouraged by the free movement of labor, generated a new migration process between countries of the Eurozone. Bandeira et al. (2019) use a DSGE model of a small open economy with heterogeneous agents in terms of their working status to analyze the consequences of fiscal austerity after the Greek economic and debt crisis. They find that labor tax hikes prompt emigration, while a reduction in government spending has a hump-shaped effect on account of a negative demand shock and a positive wealth effect associated with the expectation of lower future taxes. Other studies that introduce migration in sovereign default models (Alessandria et al., 2020) or models with search and matching frictions (Bandeira et al., 2018; Lozej, 2019) identify that the introduction of migration amplifies the cyclical fluctuations in the home economy. This amplification of the crisis is caused by changes in labor force participation (Alessandria et al., 2020; Bandeira et al., 2019; Lozej, 2019) and the reduced size of the local labor market compared to the potential number of migrants (Lozej, 2019).

The growth in international migration increased the remittance flows, with studies of the role of remittances on growth (Chatterjee and Turnovsky, 2018; Batu, 2017; Bahadir et al., 2018) and financial development (Fromentin, 2017; Aggarwal et al., 2011). Lim and Morshed (2017) stresses the importance of remittances in a two-country dynamic model with homogeneous agents to explain how fiscal policies affect labor mobility, consumption, remittances, government debt, and global output. They find that when the economy is near full employment, a tax hike on remittances can be a welfare-improving policy for the source country. Mandelman and Zlate (2012) develop a two-country DSGE model to analyze the effect of labor migration and remittance flows on smoothing consumption throughout the business cycle of the United States and Mexico, focusing on the macroeconomic consequences on the host country. Finally, Barker (2020) shows that remittances can offset the welfare losses from a negative shock on commodity prices, which led to an increase in legal and illegal migration from Venezuela to the United States.

The remaining of this document is organized as follows: Section 1.2 motivates why

Uruguay is an interesting case to study this topic. Section 1.3 introduces the model with international migration. Section 1.4 presents the data used, the calibration of the model, and its validation. Section 1.5 analyzes how changes in taxes and transfers affect incentives and outcomes. Finally, section 1.6 concludes.

1.2 Uruguay as a case study

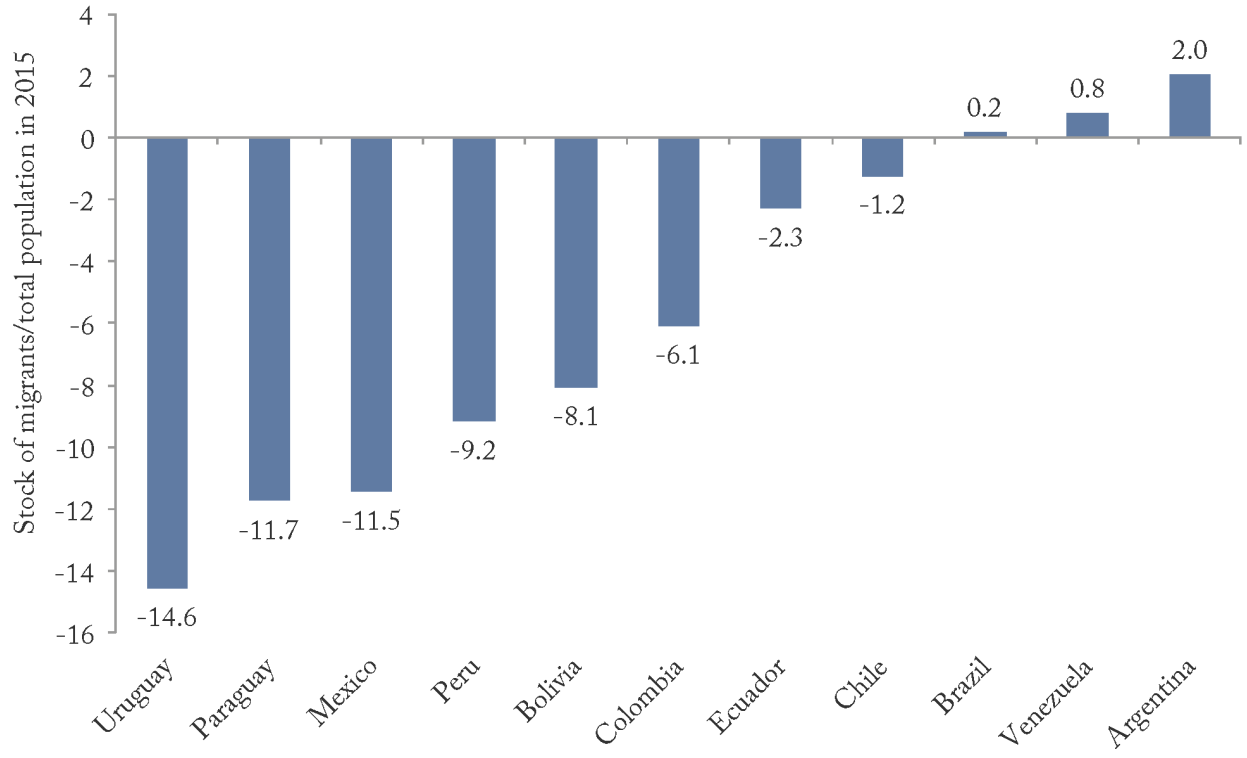
Two main reasons make Uruguay an interesting case study. The first is that, despite it being one of the smallest countries in Latin America with a total population of 3.5 million in 2017, more than 600 thousand Uruguayans (18% of its population, which is equivalent to 15% of the working-age population) live abroad (Pellegrino and Koolhaas, 2008). Figure 1.1 shows that the stock of migrants between 1950 and 2015 was almost 15% of the population of Uruguay in 2015. Meanwhile, Figure 1.2 shows that Uruguay has the lowest net migration rates among Latin American countries (including Mexico)..

International migration is a structural feature that exacerbates and reinforces the demographic aging process in Uruguay, given that 19% of the population were 60 or more years old in 2017. Migration was consolidated during the 1960s and 1970s, as it represented an alternative solution to the economic and political problems (Pellegrino, 2009). Cabella and Pellegrino (2005) estimate that between 1963 and 1975, 7% of the population emigrated from Uruguay, while between 1975 and 1985, this figure stood at around 6%. Since then, even in times of economic growth and political and social stability, net migration has been negative. Between 1999 and 2003, the country faced a huge economic and financial crisis and more than 100 thousand people (more than 3% of the resident population) left the country (Pellegrino, 2009).

In Uruguay, the volume of remittances is rather small (0.33% of the countries' GDP over the period 2002-2016, according to World Bank Data), even though a large portion of the total population lives and works abroad.[4] The low level of remittances sent back to Uruguay is related to the educational and occupational status of migrants and the relatives left behind, to the fact that people who migrate are young and in the process of emancipation from their homes, and to the level of development of the host country (Koolhaas and Pellegrino, 2009; Pellegrino and Vigorito, 2009).[5]

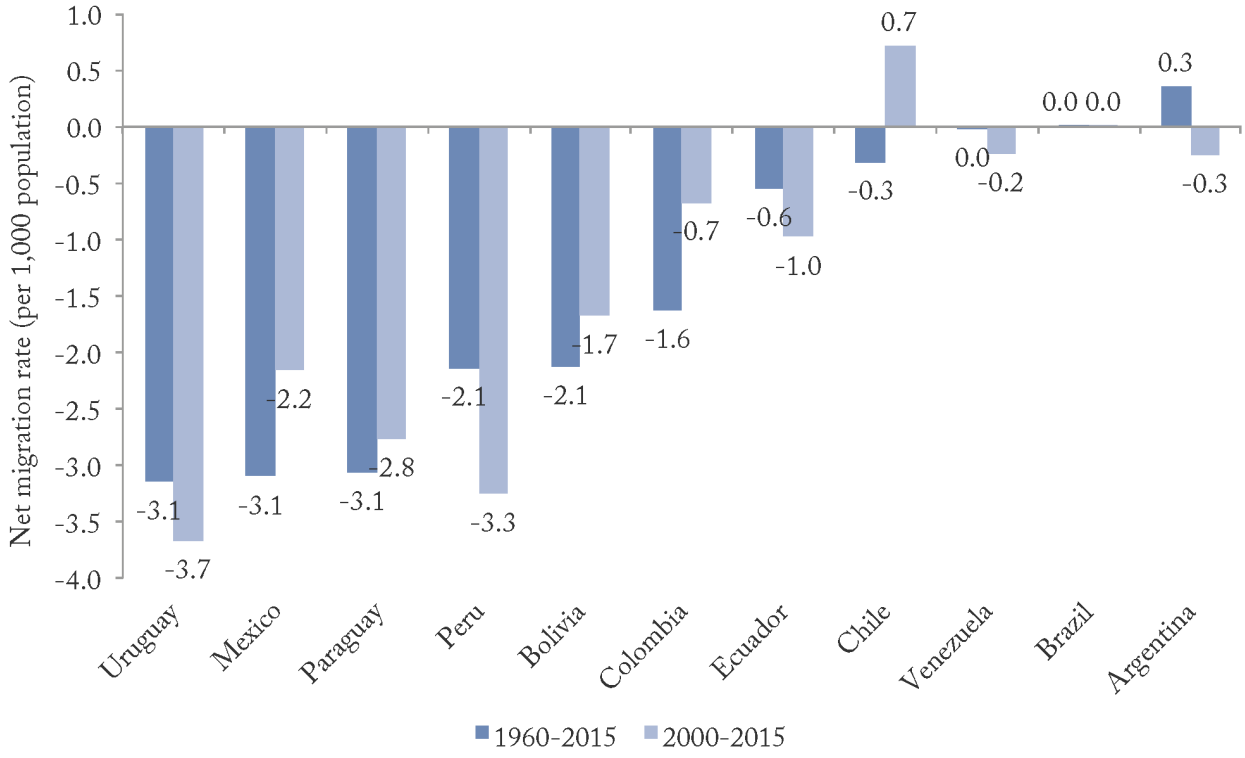
[4]Since 2002, the Central Bank of Uruguay has included family remittances among foreign transfers in its balance of payments statistics.

[5]Greece has also shown low levels of remittances in its last migration wave. Although it appears that in this case it is related to a "brain waste" mechanism, i.e., the underemployment of migrants in the host country



Source: Based on United Nations, Population Division
`http://esa.un.org/wpp/unpp/panel_population.htm`.

Figure 1.1: Stock of migrants between 1950 and 2015, over total population in 2015 (in percentage of total population). South America and Mexico



Source: Based on United Nations, Population Division
`http://esa.un.org/wpp/unpp/panel_population.htm`.
Notes: 1960-2015 denotes the fifty five-year average, and 2000-2015 the fifteen year-average.

Figure 1.2: Net migration rate per 1,000 population, 1960-2015. South America and Mexico

The second reason is that Uruguay has an extended public system of taxes and transfers. As discussed below, public social spending and the tax structure are large even for Latin American standards, with social expenditure reaching 23.5% of the GDP in the biennial average of 2011-2012; and government revenues representing 27% of the GDP in 2012. Uruguay had the highest levels of tax revenues in the 1990s and remains among the highest. Thus, the interaction of the system of taxes and transfers, and migration, is particularly relevant.

1.2.1 Public transfers

Uruguay has a long tradition of social programs that provide goods, services, as well as cash and in-kind transfers. Public social spending in Uruguay represented 23.5% of the GDP on the biennial average of 2011-2012 (CEPAL, 2016). The main component of social spending is social security, which reached 10.8% of the GDP in 2011-2012, and 90% of which was earmarked for retirements and pensions (CEPAL, 2016).

The social security system used in Uruguay was developed by the end of the nineteenth century for some specific unions, although its coverage became almost universal in the 1950s (Forteza et al., 2009). Social protection covers retirement benefits and pensions, health care, unemployment insurance, maternity benefits, disability, sickness allowances, family allowances, and non-contributory pensions for people over 70 years old. Since the social security reform in 1996, Uruguay has extended the number of years of contribution from 30 to 35, increased the minimum age for retirement, and installed a mixed pillar for contributions: one part is financed on a PAYG scheme, administrated by the (public) *Banco de Previsión Social* (BPS), and the other part is funded by individual savings accounts, operated by (private) *Administradoras de Fondos de Pensión* (AFAP). People earning below a threshold only contribute to the PAYG pillar, and people earning above it, must contribute to both. In the early 2000s, several studies showed that, under the new reform, very few people qualified to retire (Bucheli et al., 2010).

In 2008, a new reform of the social security system relaxed the number of years of contribution and the minimum retirement age. According to (Lavalleja and Tenenbaum, 2017) and (Lavalleja et al., 2018), more women can now reach retirement conditions, but it has had no effects on men. Private workers have better access and have increased the amount of their pensions; poor workers or people with fewer years of contributions can obtain a

makes it difficult to send remittances (Lazaretou, 2016).

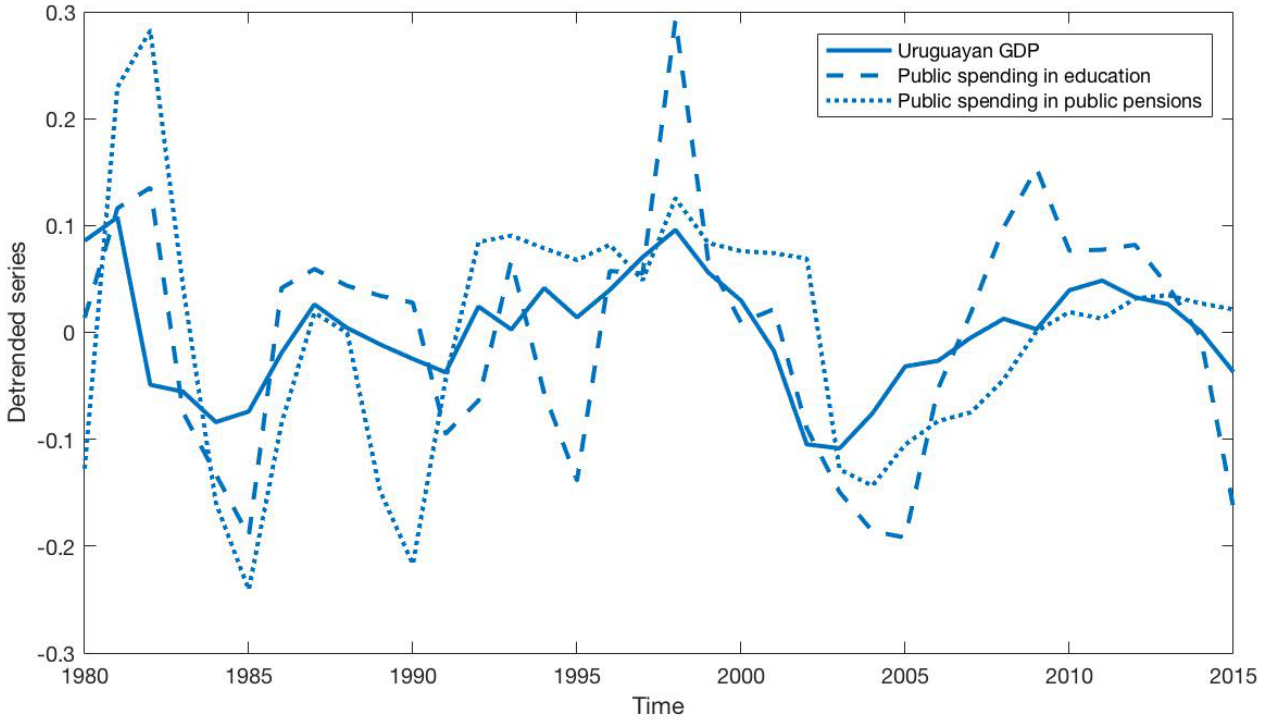
non-contributory pension but cannot access the pension for the years they had contributed. The reform also created the (public) program *Asistencia a la vejez*, which provides a subsidy to people between the ages of 65 and 69, who have no income and are in vulnerable conditions. When they turn 70, the beneficiaries receive pensions if their condition of vulnerability remains. In 2012, 83% of the population over 65 years of age received a pension from the contributory system, and 5% received it from the non-contributory system.

Regarding the public provision of health services, it began with the constitutional reform of 1934, which provided public assistance to poor people. In 2007, the public health care system was reformed and has now got two main programs: direct health care provision for poor people (non-contributory) provided by the *Administradora de Servicios de Salud del Estado (ASSE)* and the National Health System (NHS, contributory) provided by ASSE, mutual health care system, and private health care insurance. NHS is financed by a specific contribution made by employees and employers, and ASSE is financed partly by this tax and also by central government transfers. Public health care provision also covers military hospitals, police hospitals, university hospitals, and municipal polyclinics. Public health care provision coverage reached 31% of the total population in 2012 and 59% had access to NHS (INE-ECH, 2012). The NHS assists contributory workers from both the public and private sectors, as well as their children under the age of eighteen and non-contributory partners, and people from the contributory pension system. Total public spending on both programs accounted for 5.9% of the GDP in the biennial 2011-2012 (CEPAL, 2016).

Primary education in Uruguay became public, compulsory, and secular in 1877. In 2012, 83% of children in primary and middle schools attended to public schools, while 85% of high school and tertiary levels attended public institutions. In 2008 the Law of Education extended compulsory attendance up to high school, although, in practice, it is only enforced until middle school. From childcare to tertiary education (including university), Uruguayan citizens can attend different institutions of public education.[6]

As stated by Végh et al. (2017), Latin American countries differ in terms of whether they choose to apply procyclical or countercyclical fiscal policies. Uruguay is one of the countries that maintains procyclical policies. Figure 1.3 shows the cyclical components of the GDP, public spending on education, and pensions between 1980 and 2015.

[6]Although expenditures raised over the past years, problems related to the extremely high dropout and repetition rates in middle and high school have not yet been solved (Casacuberta and Bucheli, 2010; Cid and Ferrés, 2010; de Armas, 2018)



Notes: Variables expressed in natural logarithms of per-capita terms, and then detrended using the Holdrick-Prescott filter with smoothing parameter equal to 100 (annual data).

Figure 1.3: Detrended series of GDP and public spending in education and pensions. Uruguay, 1980-2015

1.2.2 Tax structure

Tax revenues have been rising in South America over the last two decades. According to the OECD, between 1990 and 2002 the value-added tax (VAT) represented over 50% of the total revenue, whereas between 2002 and 2015 the main contributors were taxes on income and profit. Unlike developed countries, which rely on direct taxes, less developed countries generate most of their revenue through indirect taxes. In 2015, 22% of the Uruguayan government's revenue came from direct taxes, 42% from indirect taxes, 28% from contributions to social security, and 8% from taxes on property (OECD, 2017).

In terms of direct taxes, since 2007, taxes on labor income and pensions are progressively taxed at rates from 0 to 30%, and capital is taxed at a flat rate of 30%. Taxes on wages and pensions represent over 52% of direct tax revenue, and corporate taxes represent 42%. In terms of indirect taxes, there are two main taxes: the VAT at a 22% tax rate, and the specific goods and services tax (IMESI) that taxes goods and services at a different rate.[7] VAT tax collection represents 70% of total indirect taxes while taxes on specific goods and services collect the remaining 30%. Direct taxes and VAT do not have a specific purpose in government spending.

The social security system is financed through direct taxes and central government transfers. Employees have a personal retirement deduction of 15% of their nominal wages, and employers contribute 7.5% of the nominal wages paid to employees. These deductions cover all of the worker's contributory benefits, excluding health care. The deduced amount for this fund changes according to income and family composition. Workers with incomes below a threshold have deductions of 3% of their nominal wage. If wages are above that threshold, deductions are of 4.5% of the nominal wage, and of 6% if the worker has dependent children under eighteen years of age. In cases where the worker chooses to include a partner in the health care system, the deduction reaches 8% of the nominal wage. Employers contribute 5% of the nominal wage of a worker covered by the NHS. Self-employed workers contribute according to a declared minimum wage.

Education in Uruguay is mainly financed by central government transfers, which are defined in the *Ley de Presupuesto Nacional* every five years. Since 1994, the *Fondo de Solidaridad* contributes to the financing of scholarships for low-income students studying at the public university of Uruguay, as well as tertiary technical careers. These transfers are financed through contributions from graduates of the public university and calculated based

[7] For example, cigarettes are taxed at 70%, some alcoholic drinks at 85%, beer at 27%, and other non-alcoholic drinks at 30%.

on the curriculum duration of the career. In turn, university graduates are required to contribute to the fund if the income generated in Uruguay exceeds a non-taxable minimum. If university graduates live abroad, they can make an application for exemption from the contribution.[8]

1.2.3 Intergenerational transfers

Children and the elderly have life-cycle deficits.[9] Children sustain their deficit through a combination of family and public transfers. The amount of the deficit funded by parents and other relatives of the family varies between countries. Public transfers at the earliest ages are education and health aimed transfers. Life-cycle deficits of the elderly also vary across countries but always tend to rely heavily on public transfers, family transfers, and asset-based reallocations.[10] Taxes and contributions fund public transfers, and private transfers rely on inter and intra-household transfers. According to Lee and Mason (2011), developing countries have a slightly increased consumption profile that contrasts with the U-shaped profile of less developed countries, and the upward profile of the more developed countries.

1.3 The model

The approximation used in this paper to investigate the relationship between the fiscal system of taxes and transfers and choices made over labor supply and migration is based on a dynamic stochastic general equilibrium model. The domestic economy is small enough to ensure there is no pressure on foreign wages when individuals emigrate; thus, only the domestic economy is modeled. There are three agents in the model: the household maximizing lifetime utility, a profit maximizing firm, and the government.

[8] According to the 2018 Annual Report of the *Fondo de Solidaridad*, 466 university graduates living abroad have received the exemption. There is no official estimate of the total number of exemptions due to this cause.

[9] As defined by Lee and Mason (2011), the life-cycle deficit is the difference between consumption and income at each age. This deficit can be funded through public or private transfers, and with public or private asset-based reallocations.

[10] Private and public asset-based reallocations could be financed with capital or natural resources, or financial assets. Capital or natural resources include land and subsoil minerals in public or private areas, as well as equipment, structures, and housing in the private sector. In the public sector, financial assets cover the debt, sovereign wealth funds, and stabilization funds. In the private sector, financial assets are in the form of consumer debt or credit, mutual funds, private pension funds and personal savings (Lee and Mason, 2011).

The household chooses consumption and labor services, both in the source and the foreign economy. The household is composed of heterogeneous individuals in age and ability, who supply units of efficiency labor to the domestic and the foreign economy.[11] The individuals who migrate send a portion of their income home after consumption. It is assumed that there is no investment (nor capital); so, households smooth their consumption through migration decisions.[12] The firm produces a homogeneous good by combining heterogeneous sources of labor. The government receives revenue from distortionary wage income taxes and consumption taxes and redistributes it through cash transfers (to the young and the elderly) and lump-sum transfers to households, balancing its budget every period. In this context, household members can leave the country to avoid excessive taxes and transfers.

1.3.1 Voluntary intergenerational transfers

1.3.1.1 Household's problem

The representative household has preferences over real consumption and labor effort, and seek to maximize the present value of the weighted utility of all members, choosing the total consumption of each household member in the home economy ($c_{s,t}$) and abroad ($c^*_{s,t}$), labor supply in the local labor market ($n_{s,t}$) and the labor supply in the foreign country ($n^*_{s,t}$). This household is composed of a unit mass of heterogeneous individuals in age and ability. Every combination of age and ability denoted by s has a mass of a_s, with $\sum_{s=1}^{S} a_s = 1$. Each household member has a unit of time that can be transformed into e_s or e^*_s efficient units of labor, where e_s is the efficiency in the local labor market and e^*_s is the efficiency in the foreign labor market. It is assumed that $e_s > 0$ and $e^*_s > 0 \; \forall s \in S$ is a constant feature in time.

Following Orsi et al. (2014), Annicchiarico and Cesaroni (2018) and Lim and Morshed

[11] Heterogeneous labor is introduced using the proposal made by Kydland (1984) and subsequently applied by Garcia-Milà et al. (2010), Maliar and Maliar (2000, 2001, 2003) and Janiak and Monteiro (2010).

[12] It is equivalent to assume that there are no capital mobility problems. Leers et al. (2004) use this assumption to simplify the analysis without losing generality.

(2017), the household maximizes the following lifetime utility:

$$\max_{\{c_{s,t},c^*_{s,t},n_{s,t},n^*_{s,t}\}^{s\in S}_{t\in T}} \mathcal{E}_0 \sum_{t=0}^{\infty} \beta^t \sum_{s=1}^{S} \psi_s a_s \left[\ln(c_{s,t}) + \delta \ln(c^*_{s,t}) - \Gamma_s \frac{(n_{s,t} + n^*_{s,t})^{1+\xi}}{1+\xi} - \Gamma^*_s \frac{(n^*_{s,t})^{1+\theta}}{1+\theta} \right] \tag{1.1}$$

where $\beta \in (0,1)$ is the inter-temporal discount factor and ψ_s is the weight of individual s in the utility of the household so that $\sum_{s=1}^{S} \psi_s = 1$. As in Lim and Morshed (2017), $\delta \ln c^*_{s,t}$ is the additively separable utility of the migrant group s from consuming in the foreign country while working abroad. The first negative term in the utility function, $\Gamma_s \geq 0$, is the disutility parameter from working in the domestic and foreign economy for agent s, while the sum of $n_{s,t}$ and $n^*_{s,t}$ represents the aggregate participation rate of group s. The second negative term represents the specific cost of working abroad, which is identified with the parameter $\Gamma^*_s \geq 0$.[13] The introduction of an additional disutility on foreign labor supply is due to the subjective cost of leaving family or friends, while working abroad.[14] The parameters ξ and θ are the inverse of the Frisch elasticities of aggregate and foreign labor supplies.

The household has net incomes from working in the domestic economy, $w_t \sum_{s=1}^{S} (1-\tau^s_w) a_s e_t n_{s,t}$, where τ^s_w is the labor income tax rate to group s; incomes from working abroad, $w^*_t \sum_{s=1}^{S} a_s e^*_s n^*_{s,t}$; and revenues that come from the profit maximization problem of the firm, Π_t. The government carries out transfers that fund inactivity between young (κ^y_s) and senior (κ^e_s) citizens: young people use this transfer to study while elders use it for retirement. Total in-cash transfers are given by $\kappa^y_t \sum_{s=1}^{S} a_s \zeta^y_s \ell_s + \kappa^e_t \sum_{s=1}^{S} a_s \zeta^e_s \ell_s$, where $\ell_{s,t}$ is the leisure of group s $\left(\ell_{s,t} = 1 - n_{s,t} - n^*_{s,t}\right)$. Finally, the household receives a lump-sum transfer (T_t).

[13] This specification allows for labor adjustments along the intensive and extensive margins. The extensive margin takes into account the decision of emigrating or staying at home.

[14] Busato and Chiarini (2004), Busato et al. (2012), Orsi et al. (2014) and Annicchiarico and Cesaroni (2018) have interpreted this additional cost as capturing the lack of social and health insurance in the underground sector.

Therefore, the household maximization problem faces the following budget constraint:

$$w_t \sum_{s=1}^{S} (1-\tau_w^s) a_s e_t n_{s,t} + w_t^* \sum_{s=1}^{S} a_s e_s^* n_{s,t}^* + \Pi_t + \kappa_t^y \sum_{s=1}^{S} a_s \zeta_s^y \ell_s + \kappa_s^e \sum_{s=1}^{S} a_s \zeta_s^e \ell_s + T_t \geq \tag{1.2}$$

$$(1+\tau_c) \sum_{s=1}^{S} a_s c_{s,t} + \sum_{s=1}^{S} a_s c_{s,t}^* + \sum_{s=1}^{S} a_s \frac{\phi_s}{2} (n_{s,t}^* - n_{s,t-1}^*)^2 \tag{1.3}$$

Income finances consumption of all family members in the domestic economy $\left((1+\tau_c)\sum_{s=1}^{S} a_s c_{s,t}\right)$, abroad $\left(\sum_{s=1}^{S} a_s c_{s,t}^*\right)$ and quadratic migration costs $\left(\sum_{s=1}^{S} a_s \frac{\phi_s}{2}(n_{s,t}^* - n_{s,t-1}^*)^2\right)$. τ_c is a distortionary tax on consumption. Migration costs are added to smooth out-migration over the business cycle and other shocks, through to the absence of assets or savings in the model.[15]

Migrants send a fraction of their labor income home every period (ι), after deciding on consumption. Thus, consumption abroad is expressed as:

$$\sum_{s=1}^{S} a_s c_{s,t}^* = (1-\iota) w_t^* \sum_{s=1}^{S} a_s e_s^* n_{s,t}^* \tag{1.4}$$

and remittances are defined as:

$$R_t = \iota w_t^* \sum_{s=1}^{S} a_s e_s^* n_{s,t}^* \tag{1.5}$$

The local economy is assumed to be small. Thus, the wage earned abroad is exogenous and it follows the law of motion:

$$\ln(w_t^*) = (1-\rho_{w^*})\bar{w}^* + \rho_{w^*} \ln(w_{t-1}^*) + v_t \tag{1.6}$$

where $\bar{w}^*$ is the long-term average of the wage in the foreign country and v is an i.i.d. normal error term, $v_t \sim \mathcal{N}(0, \sigma_v^2)$.

[15] Apart from its mathematical convenience, quadratic costs are relatively easy to calibrate as we treat the costs of migrating or returning symmetrically.

1.3.2 The firm

The representative firm produces a homogeneous good used for consumption. The technology is a function of local heterogeneous labor with constant returns to scale. Thus, the profit maximization problem of the firm is given by:

$$\begin{aligned} \max_{\{n_{s,t}\}} \ \Pi_t &= y_t - w_t h_t \\ \text{subject to } \ y_t &= A \exp(z_t) h_t^{\alpha} \ \ \forall t \end{aligned} \tag{1.7}$$

where α is the share of labor in total domestic production, as well as the elasticity of the demand curve of the firm, $h_t \equiv \sum_{s=1}^{S} a_s e_t n_{s,t}$ is the labor measured in efficiency units of labor services in period t, and w_t is the wage paid for one unit of efficiency labor in the domestic economy. z_t is an aggregate productivity stochastic shock, which follows an $AR(1)$ process with independent and identically distributed normal error term:

$$\ln(z_t) = \rho_z \ln(z_{t-1}) + \varepsilon_t \tag{1.8}$$

with $0 < \rho_z < 1$ and $\varepsilon_t \sim \mathcal{N}(0, \sigma_\varepsilon^2)$. The first order condition for the firm implies:[16]

$$w_t = \alpha A \exp(z_t) h_t^{\alpha - 1} \tag{1.9}$$

The firm demands efficiency units of labor until its marginal productivity equals its marginal cost.

1.3.3 The government

The government taxes consumption and local labor income and carries out transfers to the young and the elderly. The transfers to young people are to fund their education, while the transfers to the elderly are to fund their retirement. The government redistributes its revenue in the local economy and balances its budget each period, so that:

$$\kappa_t^y \sum_{s=1}^{S} a_s \zeta_s^y \ell_s + \kappa_t^e \sum_{s=1}^{S} a_s \zeta_s^e \ell_s + T_t = \tau_c \sum_{s=1}^{S} a_s c_{s,t} + w_t \sum_{s=1}^{S} \tau_w^s a_s e_s n_{s,t} \tag{1.10}$$

[16] Since the economy only produces a homogeneous consumption good, the equilibrium price vector has been normalized to (1,w_t), where w_t denotes the equilibrium real wage rate.

where κ_t^i is the total transfer to age group i, with $i = y, e$ representing young and elder people, respectively; ζ_s^i is the ratio of in-cash transfers i given to age group s and $\sum_{s=1}^{S} \zeta_s^i = 1$.[17] Iqbal and Turnovsky (2008) introduced an alternative way to model intergenerational government transfers by giving young people and the elderly a different fraction of total output. As in their work, the proposal in this paper allows for transfers to generate potential intra- and intertemporal distortions on the labor supply decisions of all family members. As the correlation between government transfers and output in the data is not perfect, Arseneau and Chugh (2012) are followed, therefore assuming that government transfers to the young and elders follow the exogenous process given by:[18]

$$\ln(\kappa_t^i) = (1 - \rho_{\kappa^i})\bar{\kappa}^i + \rho_{\kappa^i} \ln(\kappa_{t-1}^i) + u_t^i \tag{1.11}$$

where $\bar{\kappa}^i$ is the long-term value of transfers i, u_t^i is an idiosyncratic shock to government transfers i in period t with $u^i \sim \mathcal{N}(0, \sigma_{u^i}^2)$. These shocks are independent from each other and identically distributed, but they are all correlated to the technology shock. It is also assumed that there is no correlation between the foreign wage shock and the home stochastic processes.

1.3.4 Competitive equilibrium

The market clearing condition for the domestic labor market is:

$$h_t^D = \sum_{s=1}^{S} a_s e_s n_{s,t} \quad \forall t \tag{1.12}$$

At the aggregate level, the resource constraint (feasibility condition) is given by:

$$\sum_{s=1}^{S} a_s c_{s,t} + \sum_{s=1}^{S} a_s \frac{\phi_s}{2} (n_{s,t}^* - n_{s,t-1}^*)^2 = y_t + \iota w_t^* \sum_{s=1}^{S} a_s e_s^* n_{s,t}^* \tag{1.13}$$

In the domestic economy, the sum of consumption and migration costs must equal the domestic product and the remittances sent home by migrants.

Definition 1 *A competitive equilibrium consists of a set of weights for* $\{\psi_s\}^{s \in S}$*, a sequence of*

[17] For the application below, young people are defined as those under 25 years old, and elders as those over 59 years old. Other age groups do not have in-cash transfers.

[18] Arseneau and Chugh (2012) assume an exogenous process for total government spending.

contingency plans $\{\{c_{s,t}, c^*_{s,t}, n_{s,t}, n^*_{s,t},\}^{s\in S}\}_{t=0}^{\infty}$, $\{h_t\}_{t=0}^{\infty}$, *the equilibrium domestic wage* $\{w_t\}_{t=0}^{\infty}$ *and the laws of motion of the exogenous processes,* $\{w^*_t, z_t, \kappa^y_t, \kappa^e_t\}_{t=0}^{\infty}$ *such that:*

- *Given* $\{w_t\}_{t=0}^{\infty}$ *and* $\{w^*_t, z_t, \kappa^y_t, \kappa^e_t\}_{t=0}^{\infty}$, $\{\{c_{s,t}, c^*_{s,t}, n_{s,t}, n^*_{s,t},\}^{s\in S}\}_{t=0}^{\infty}$ *solve the household problem (1.1) subject to (1.2).*
- *Given* $\{w_t\}_{t=0}^{\infty}$ *and* $\{w^*_t, z_t, \kappa^y_t, \kappa^e_t\}_{t=0}^{\infty}$, $\{h_t\}_{t=0}^{\infty}$ *solve the firm problem (1.7).*
- *The government satisfies (1.10).*
- *Markets clear. That is, equations (1.12) and (1.13) are satisfied.*

The first order optimality conditions for this economy are:

$$c_{s,t} : \psi_s c_{s,t}^{-1} - \lambda_t(1+\tau_c) \leq 0 \quad \forall s,t \tag{1.14}$$

$$c^*_{s,t} : \psi_s \delta c^{*^{-1}}_{s,t} - \lambda_t \leq 0 \quad \forall s,t \tag{1.15}$$

$$n_{s,t} : -\psi_s \Gamma_s (n_{s,t} + n^*_{s,t})^{\xi} + \lambda_t \left[(1-\tau^s_w) w_t e_t - \kappa^i_t \zeta_s\right] \leq 0 \quad \forall s,t \tag{1.16}$$

$$\begin{aligned} n^*_{s,t} : &-\psi_s \left[\Gamma_s (n_{s,t} + n^*_{s,t})^{\xi} + \Gamma^*_s n^{*\theta}_{s,t}\right] + \\ &\lambda_t \left[w^*_t e^*_s - \phi_s (n^*_{s,t} - n^*_{s,t-1}) - \kappa^i_t \zeta_s\right] + \beta \phi_s \mathcal{E}_t \lambda_{t+1} (n^*_{s,t+1} - n^*_{s,t}) \leq 0 \quad \forall s,t. \end{aligned} \tag{1.17}$$

For each period t, household chooses each member's consumption level at home and abroad, as well as the labor supply at home and abroad, taking into account that these decisions have intertemporal consequences. Quadratic migration costs ensure that previous migration decisions are taken into account in the optimization problem ($n^*_{s,t-1}$), and lead to a nonlinear second-order differential equation.

Establishing λ_t as the marginal utility of real income, from the first-order conditions (1.16) and (1.17), we obtain:

$$n_{s,t} + n^*_{s,t} = \left\{ \frac{\lambda_t}{\psi_s \Gamma_s} [(1-\tau^s_w) w_t e_t - \kappa^i_t \zeta_s] \right\}^{\frac{1}{\xi}} \tag{1.18}$$

$$\begin{aligned} n^*_{s,t} = \Bigg\{ &\frac{\lambda_t}{\psi_s \Gamma^*_s} [w^*_t e^*_s - (1-\tau^s_w) w_t e_t - \phi_s (n^*_{s,t} - n^*_{s,t-1})] + \\ &\frac{\phi_s}{\psi_s \Gamma^*_s} \beta \mathcal{E}_t \lambda_{t+1} (n^*_{s,t+1} - n^*_{s,t}) \Bigg\}^{\frac{1}{\theta}} \end{aligned} \tag{1.19}$$

Equation (1.18) defines total labor supply as a positive function of the real wage of the local labor market minus government transfers, weighted by the marginal utility of con-

sumption and the disutility from overall work. In turn, (1.19) defines how the household chooses the migration rate optimally. The household will supply labor abroad if the wage rate offsets the wage differential and migration costs of the current period and the next one.

In this specification, ξ determines the responsiveness of total labor supply to an increase in net real earnings in the local labor market, while θ determines the responsiveness of migrant labor supply to the wage differential and migration costs. Lower levels of ξ or θ lead to higher volatility in total and foreign labor supply, respectively.

Given the structure of the economy and the way in which households fund their consumption, changes in these parameters can affect the labor supply from the family members directly involved as well as the remaining members of the household. The order of magnitude depends on Frisch's elasticity of substitution, the disutility of work for each member, their abilities in the domestic and foreign economies, and the weight of each member on total utility.

1.4 Migration and life cycle profiles

Migration is not random. Age, education, ability to face risks, and adapt to new contexts are some of the characteristics that determine the migration profile of individuals. Heterogeneity in age and ability is introduced to address some of these aspects. Five age groups and two types of labor (skilled and unskilled) are defined. The age groups are set to represent life cycle profiles: those under the age of 25, 25 to 34 years old, 35 to 49 years old, 50 to 59 years old, and older than 60 years old. Meanwhile, the ability groups are set to analyze the response of individuals to reforms in the tax and transfer system when there are no restrictions to migrate. Unskilled people are those who have achieved less than 12 years of formal education (high-school level), while skilled people have had 12 years or more of formal education.[19]

1.4.1 Data

Labor supply, wages, and consumption are estimated based on cross-sectional data from the Uruguayan economy and Uruguayan migrants living in the United States, Argentina, and Spain. Yearly cross-sectional data of Uruguay covers from 1982 to 2016. Steady-state values for local variables of the labor market are estimated using the Encuesta Continua

[19]For people under 20 years of age, the skill classification is established for the head of the household.

de Hogares from the Instituto Nacional de Estadística (ECH-INE) of Uruguay of 2006. The choice of the year 2006 is due to a range of reforms in the Uruguayan tax and transfer system that started in 2007. Data series of GDP and consumption are taken from the Banco Central del Uruguay (BCU), and consumption profiles are estimated using data from the National Transfer Accounts of Uruguay.[20] The total population information comes from INE, while the number of migrants is collected from Cabella and Pellegrino (2005) and supplemented with information brought from the Dirección Nacional de Emigración in Uruguay. Data on remittances are from the Annual Remittances Data (World Bank `https://www.worldbank.org/en/topic/migrationremittancesdiasporaissues/brief/migration-remittances-data`).

Labor and earning profiles for Uruguayans living abroad are estimated considering the distribution of emigrants. As shown in Table 1.1, in 2008 more than 30% of Uruguayans living abroad were in Argentina, 17% lived in the United States of America, 15% in Spain, and 15% in Brazil. The *Encuesta Permanente de Hogares* database from Argentina between 2003 and 2016, the American Community Survey from the United States between 2000 and 2016, and the immigrants survey from the National Institute of Statistics of Spain for 2007 were used.[21] Table 1.2 presents the distribution of Uruguayans living in Uruguay and abroad, according to age and ability. The proportion of Uruguayans living abroad is increasing in the age group, and among those in the high ability group.

[20]Since there is no estimation of the consumption profiles of Uruguayan migrants, it is assumed that they are the same as the ones observed for Uruguayans living in Uruguay.

[21]There is no micro-data available for Uruguayans living in Brazil.

Table 1.1: Distribution of Uruguayan population, by country of residence (in %)

	World distribution		Percentage of Uruguayans by country of residence	
	2008	2016	2008	2016
Uruguay	85%	87%		
Argentina	5%	3%	31%	27%
United States of America	3%	2%	17%	16%
Spain	2%	2%	15%	18%
Brazil	2%	2%	15%	13%
Other Latin American countries	1%	1%	7%	10%
Rest of the world	2%	2%	15%	16%

Notes: Unofficially, it is estimated that there were more than 213.650 and 212.800 Uruguayans living in Argentina in 2008 and 2016, respectively.
Source: Ministerio de Relaciones Exteriores, Dirección General para Asuntos Consulares y Vinculación del Uruguay.

Table 1.2: Distribution of Uruguayan population, by age and ability groups (in %), 2006

Age group	Ability group	Uruguayan population	Residence (in %)	
		(in millions)	Uruguay	Abroad
0-25	Unskilled	1.08	0.95	0.05
	Skilled	0.28	0.90	0.10
25-34	Unskilled	0.37	0.83	0.17
	Skilled	0.17	0.73	0.27
35-49	Unskilled	0.55	0.82	0.18
	Skilled	0.25	0.66	0.34
50-59	Unskilled	0.33	0.85	0.15
	Skilled	0.13	0.66	0.34
60 and +	Unskilled	0.64	0.89	0.11
	Skilled	0.13	0.67	0.33

Notes:Calculations based on data of the Ministerio de Relaciones Exteriores, Dirección General para Asuntos Consulares y Vinculación del Uruguay, Encuesta Continua de Hogares (INE, Uruguay), Encuesta Permanente de Hogares (INDEC, Argentina), American Community Survey (U.S. Census Bureau) and Encuesta Nacional de Emigrantes de España (INE, España).

1.4.2 Calibration

The model is calibrated using annual data to match the average steady-state of the Uruguayan economy in the year 2006 and data of Uruguayans living abroad.

The discount factor (β) is set to 0.95, consistent with international DSGE models, and the contribution of labor to total output (α) is set to 0.67, in line with the estimated social contribution matrix for Uruguay (Katz et al., 2004). Table 1.3 summarizes the parametrization done to this economy.

Table 1.3: Parametrization.

Parameter	Description	Value
Calibrated		
β	Discount factor	0.95
α	Work participation in total output	0.67
ξ	Inverse of elasticity of total labor supply	4.00
θ	Inverse of elasticity of foreign labor supply to wage gap	0.30
τ_c	Consumption implicit tax rate	0.10
$\bar{w}^*$	Long term average of the wage in foreign country	4.00
Implied values		
A	Total factor productivity	2.00
ι	Share of incomes earned abroad sent to home	0.01
δ	Weight on migrants utility from consumption in foreign	0.28
κ^y	Government transfers to young people (education)	0.34
κ^e	Government transfers to elders (retirement)	1.46

To calibrate the inverse of the elasticity of aggregate labor supply (ξ) and the inverse of the elasticity of foreign labor supply (θ), the wage differentials are calibrated following the suggestions made by Peterman (2016). Particularly, the calibration matches the first two moments of aggregate labor supply at home and abroad, and the implied volatility of migration flows. The estimation of the intensive margin is 4 and that of the extensive margin is 0.3.[22] The implicit consumption tax rate is set to 10%, which reproduces the effective revenue from the VAT. The long term average wage abroad is set to 4 ($\bar{w}^*$).

Table 1.4 presents the calibrated parameters for the members of the household. With this parameterization, employment at home and abroad is obtained, as well as the wage

[22] See Annicchiarico and Cesaroni (2018) for details.

profile for each household member.[23] The mass of individuals is the relative weight of each combination of age and ability in the total population. The weight of the consumption of each group in the total utility of the family is computed with the estimation made by Bucheli and González (2011) of total consumption by age and educational groups for Uruguay in 2006. The household's labor income implicit tax rate is defined as the share of estimated taxes paid in each age-ability group, relative to their nominal wage. This tax includes taxes on labor income, social security contributions and health care insurance. The efficiency parameters of the foreign and the domestic economy are calculated as the ratio between the hourly wage of each group and the hourly wage of the foreign and the local economy, respectively. Finally, the share of in-cash transfers to the young and the elderly match the share of total transfers on the GDP of each age-ability group in Uruguay for 2006.

[23]The model does not have unemployment. Thus, employment in the home economy (abroad) is defined as the ratio between the employed in the home economy (abroad) and the total employed (at home and abroad) plus the population outside the workforce (home and abroad).

Table 1.4: Family parameterization.

	Calibrated						Implied values	
Parameter	a_s [a]	ψ_s [b]	τ_w^s [c]	$\bar{e_s}$ [d]	e_s^* [e]	ζ_s [f]	Γ_s [g]	Γ_s^* [h]
Unskilled								
0 to 24 years old	0.27	0.04	0.10	0.50	0.72	0.58	7.89	7.06
25 to 34 years old	0.08	0.07	0.12	0.66	0.78	0.00	2.00	3.42
35 to 49 years old	0.16	0.07	0.12	0.82	0.82	0.00	1.88	3.11
50 to 59 years old	0.08	0.07	0.11	0.86	0.83	0.00	2.87	3.44
60 years and older	0.17	0.08	0.07	0.77	0.78	0.32	50.67	3.71
Skilled								
0 to 24 years old	0.07	0.11	0.15	0.88	0.85	0.42	8.57	2.52
25 to 34 years old	0.03	0.13	0.18	1.27	1.04	0.00	1.65	1.74
35 to 49 years old	0.07	0.13	0.19	1.88	1.22	0.00	1.64	1.58
50 to 59 years old	0.04	0.14	0.18	2.21	1.21	0.00	2.24	1.01
60 years and older	0.03	0.16	0.14	2.11	1.01	0.68	23.45	0.61

Notes:
[a] Share of population.
[b] Weight in the total utility of the family.
[c] Household labor income implicit tax rate.
[d] Efficiency units of labor services in home economy.
[e] Efficiency units of labor services in the foreign economy.
[f] Share of government in-cash transfers.
[g] Disutility of aggregate labor.
[h] Disutility of labor supply in foreign economy.

This calibration determines some of the values of the parameters of this model. For example, in Table 1.3, after the long-term average wage in the foreign country is determined, the total factor productivity in the home economy is calibrated to fit the log wage differential. Similarly, the share of income earned abroad and sent home is established to match remittances as a share of output in the home economy (0.5% of GDP); the weight of the migrant's utility from consumption is set to match foreign consumption after setting the share of remittances relative to GDP, and government transfers to young people and the elderly are calibrated to match the share of transfers to each age-group in the GDP.[24] Table 1.4 also presents the implied values associated with disutilities from aggregate work and from working abroad. They are calibrated to match the aggregate employment rate and the rate

[24] As noted by the reviewer, the share of income earned abroad and sent home is low because the model assumes that every migrant remit, and at the aggregate, must reproduce the large portion of the population living abroad and the share of remittances relative to the GDP that Uruguay receives.

of employment abroad. The disutility of total work is similar between skilled and unskilled people and mimics the U-shape of labor supply. Parametrization of disutility from working abroad is higher among unskilled workers, and for both types of labor, it is decreasing within age groups. This result may be explained by higher implicit costs in individuals with less ability.

The model matches the steady-state values of the main macroeconomic variables of the model, summarized in Tables 1.5 and 1.6.[25] Employment rates at home and abroad replicate estimated life cycle profiles. These profiles have an inverted U-shape, in line with life cycle estimations.

Table 1.5: Deterministic steady-state of the main variables.

Variable	Description	Value
y	Output	1.16
w	Wage at home	1.75
$\sum_{s=1}^{S} a_s e_s n_s$	Total employment at home	0.47
$\sum_{s=1}^{S} a_s e_s^* n_s^*$	Total employment at foreign	0.12
$\sum_{s=1}^{S} a_s c_s$	Total consumption at home	1.17
$\sum_{s=1}^{S} a_s c_s^*$	Total consumption at foreign	0.43
	Stylized facts	
$\ln\left(\frac{w^*}{w}\right)$	Pre-tax log wage difference (in %)	82.5%
R/y	Remittances-GDP ratio	0.5%
G_y	Government transfers to young-GDP ratio (in %)	2.9%
G_e	Government transfers to elders-GDP ratio (in %)	7.0%

Notes: Government transfers to elders include pensions to people aged 60 and over. Total pensions (including other age groups) reached 8.3% of the Uruguayan GDP in 2006.

[25] Table 1.5 shows two relevant issues of Uruguayans living abroad. Firstly, people who leave Uruguay earn, on average, a wage measured in PPP that is 2.3 times higher than the average wage earned in this country. Secondly, they remit home a low proportion of incomes generated abroad, representing only 0.5% of the GDP in the year 2006. Figure 1.A.1 from the Appendix shows the mean pre-tax wage difference between wages earned abroad and home, for the year 2006, for each age and skilled group.

Table 1.6: Deterministic steady-state of employment rates in the home and foreign economies by age and ability groups.

Variable	Employment at home (n_s)[a]	Employment at foreign (n_s^*)[b]
Unskilled		
0 to 24 years old	0.42	0.06
25 to 34 years old	0.63	0.16
35 to 49 years old	0.68	0.16
50 to 59 years old	0.65	0.08
60 years and older	0.19	0.04
Skilled		
0 to 24 years old	0.35	0.07
25 to 34 years old	0.56	0.25
35 to 49 years old	0.64	0.28
50 to 59 years old	0.55	0.30
60 years and older	0.19	0.14

Notes: Labor supply in Uruguay (abroad) is determined by the ratio of people employed in Uruguay (abroad) over the sum of people working and those over 16 years of age out of the labor force in Uruguay (abroad).

[a] $n_s = \dfrac{E_{s,t}}{E_{s,t} + olf_{s,t} + E^*_{s,t} + olf^*_{s,t}}$

[b] $n^*_s = \dfrac{E^*_{s,t}}{E_{s,t} + olf_{s,t} + E^*_{s,t} + olf^*_{s,t}}$

1.4.3 Estimation

The perturbation method described by Schmitt-Grohé and Uribe (2004) was used to solve the dynamic model. This method applies a second-order approximation to the model equilibrium around the deterministic steady-state. The equilibrium conditions to solve the model correspond to equations (1.6), (1.8), (1.9), (1.11), and (1.14)-(1.17), for all $s \in S$.

The parameters that conduct the dynamics of the model are estimated to match the first two moments of output, wage abroad, government transfers for education and pensions, migration, and the correlation between output and government transfers. The estimation of the parameters is made using the method of moments, and is presented in Table 1.7. The target moments (presented in Table 1.8) are the autocorrelations of domestic output (0.686), foreign wage (0.629), transfers to education (0.521) and transfers to pensions (0.571); volatility of output (5.06), volatility of foreign wage (4.72), volatility of education (11.0) and

pensions (11.38); and the cyclical correlation between output and education and pension transfers (0.668 and 0.451, respectively). Finally, using data from migration flows, the costs of migration were calibrated to match the volatility of migration in each age-ability group.

Table 1.7: Summary of estimated parameters.

Variable	Description	Value
ρ_z	Autocorrelation of technology shock	0.515
σ_z	Standard deviation of technology shock	0.032
ρ_{w*}	Autocorrelation of foreign wage	0.628
σ_{w*}	Standard deviation of foreign wage	0.037
ρ_{κ^y}	Autocorrelation of education shock	0.515
σ_{κ^y}	Standard deviation of education shock	0.094
$corr_{z,\kappa^y}$	Correlation between technology and education shock	0.768
ρ_{κ^e}	Autocorrelation of pensions shock	0.569
σ_{κ^e}	Standard deviation of pensions shock	0.093
$corr_{z,\kappa^e}$	Correlation between technology and pensions shock	0.471
Adjustment costs of migration		
Unskilled		
$\phi_{U,0-24}$	0 to 24 years old	8.53
$\phi_{U,25-34}$	25 to 34 years old	20.14
$\phi_{U,35-49}$	35 to 49 years old	18.51
$\phi_{U,50-59}$	50 to 59 years old	26.56
$\phi_{U,60+}$	60 years and older	19.81
Skilled		
$\phi_{S,0-24}$	0 to 24 years old	15.47
$\phi_{S,25-34}$	25 to 34 years old	33.94
$\phi_{S,35-49}$	35 to 49 years old	25.00
$\phi_{S,50-59}$	50 to 59 years old	19.54
$\phi_{S,60+}$	60 years and older	16.83

Table 1.8 reports the simulated moments from the calibrated benchmark model and contrasts them with the empirical estimation. The upper part of the table shows the standard deviation, the first-order autocorrelation, and correlation with GDP for output, foreign wage and government transfers for education and pensions.[26] Additionally, the second part of the table presents untargeted second moments of the model: home consumption, aggregate employment rate at domestic and foreign economies, and wage at home. As can

[26]Figure 1.A.2 present the simulated series of the main aggregate variables of the model.

be noted, the sign of the simulated series is in line with the data, and some magnitudes are very similar, indicating that the model captures these moments accurately even though they are untargeted. Table 1.9 presents the targeted second moments of migration: this is the volatility of each age-ability group.

Table 1.8: Business cycle moments.

	Data	Model Benchmark case
Targeted second moments		
σ_{y_t}	5.06 (4.20;6.35)	5.06
ρ_{y_t}	0.686 (0.52;0.85)	0.686
$\sigma_{w_t^*}$	4.72 (4.45;7.44)	4.72
$\rho_{w_t^*}$	0.629 (0.03;1.23)	0.629
$\sigma_{\kappa_t^y}$	11.0 (9.22;14.84)	11.0
$\rho_{\kappa_t^y}$	0.521 (0.23;0.81)	0.521
$corr(\kappa_t^y, y_t)$	0.668 (0.44;0.82)	0.668
$\sigma_{\kappa_t^e}$	11.38 (8.91;14.34)	11.38
$\rho_{\kappa_t^e}$	0.571 (0.33;0.81)	0.571
$corr(\kappa_t^e, y_t)$	0.451 (0.15;0.68)	0.451
Untargeted second moments		
$corr(w_t^*, y_t)$	-0.108 (-0.58;0.42)	-0.325
σ_{C_t}	6.03 (5.00;7.57)	4.97
ρ_{C_t}	0.741 (0.54;0.94)	0.681
$corr(C_t, y_t)$	0.927 (0.87;0.96)	0.999
σ_{N_t}	1.18 (0.95;1.55)	1.55
ρ_{N_t}	0.520 (0.17;0.87)	0.928
$corr(N_t, y_t)$	0.687 (0.46;0.83)	0.687
$\sigma_{N_t^*}$	0.68 (0.49;1.07)	1.96
$\rho_{N_t^*}$	0.503 (0.14;0.87)	0.912
$corr(N_t^*, y_t)$	-0.311 (-0.71;0.24)	-0.781
σ_{w_t}	9.24 (7.45;12.15)	3.68
ρ_{w_t}	0.643 (0.39;0.9)	0.496
$corr(w_t, y_t)$	0.530 (0.24;0.73)	0.708

Notes: Capital letters correspond to aggregate variables. Estimations using 12.000 periods, dismissing the first 2.000. 95% confidence intervals in parenthesis.

Table 1.9: Business cycle moments: migration by age-ability group.

	Data	Model (Benchmark)
Targeted second moments Standard deviation of labor migration flows $(\sigma_{n^*_{s,t}-n^*_{s,t-1}})$		
Unskilled		
0 to 24 years old	0.83	0.83
25 to 34 years old	0.99	0.99
35 to 49 years old	1.11	1.11
50 to 59 years old	0.69	0.69
60 years and older	0.58	0.58
Skilled		
0 to 24 years old	0.85	0.85
25 to 34 years old	1.02	1.02
35 to 49 years old	1.47	1.47
50 to 59 years old	1.81	1.81
60 years and older	1.53	1.53

Notes: Estimations using 12.000 periods, dismissing the first 2.000.

1.5 Results

1.5.1 Temporary shocks in the intergenerational transfer system

The effect of an exogenous shock of 1% in educational transfers and pensions transfers on total labor supply, labor supply at home and abroad are presented in Figures 1.4 and 1.5, respectively. As can be expected from equation (1.18), the members of the family that receive the transfer reduce their total labor supply. The negative income effect derived from the lower participation of the affected groups propels other age-ability groups to increase their total labor supply. Equation (1.19) defines the optimal migration rate of each group depending on a variety of parameters. As a result, changes in intra- and intertemporal labor supply at home and migration occur along different age-ability groups.[27]

Figure 1.4 shows that an unexpected educational transfer shock has a higher effect on the younger members of the family, who reduce their total labor supply, particularly their labor

[27] Additional exercises of impulse-response functions are available upon request.

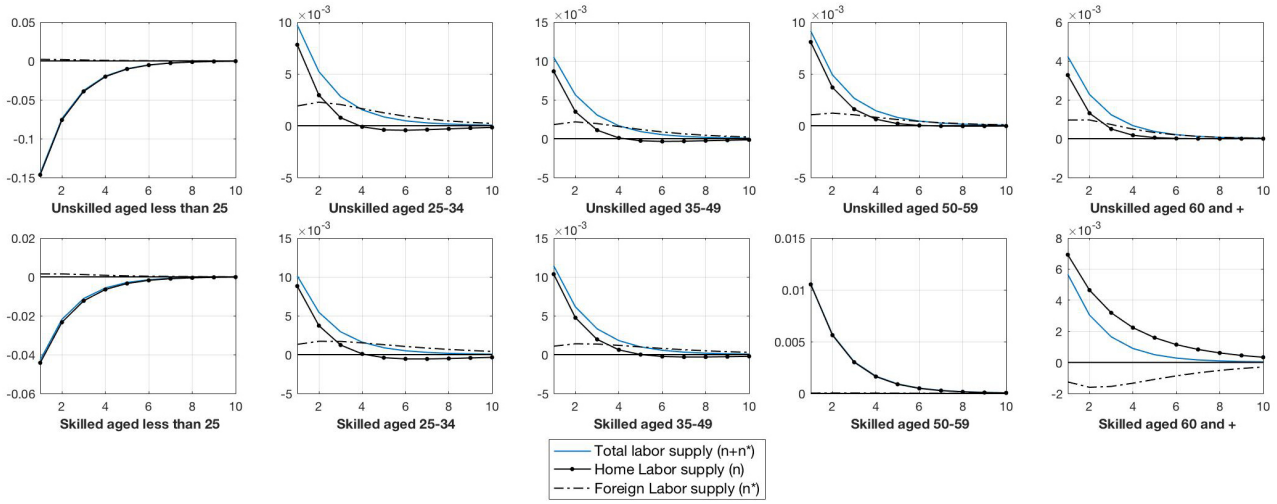
supply at home. Moreover, the educational transfer slightly increases their participation rate abroad. This effect is not obvious and occurs because young people earn relatively more working abroad and face lower migration costs, making migration more attractive to smooth consumption at the family level. The skilled and unskilled workers aged between 25 and 60 increase their total labor supply (at home and abroad) to compensate for the negative income effect that results from the lower labor supply at home from the young. Adults over 60 years of age increase their total labor supply, but the skilled and unskilled groups react differently regarding their labor supply abroad. Skilled workers over 60 years of age have an incentive to work at home, as it generates higher incomes than abroad, while the unskilled group of people over 60 years of age is more efficient working abroad. Since they both have to increase their total labor supply, the optimal response for each group is to change their migration decisions in opposite directions.

In the case of an unexpected pension shock (Figure 1.5), total labor supply of the elderly is reduced, mainly driven by the reduction in the labor supply at home. There are minor movements in the labor supply abroad from this group. As before, the skilled group of people comes back home to take advantage of the shock, but the unskilled group faces a moderate increase in the labor supply abroad. The skilled group is more productive at home and the unskilled abroad, and it is optimal for the family to choose a different path on foreign labor supply. Once again, the negative income effect propels the labor supply in the home economy for groups not directly affected by this shock.

1.5.2 Permanent changes in the intergenerational tax and transfer system

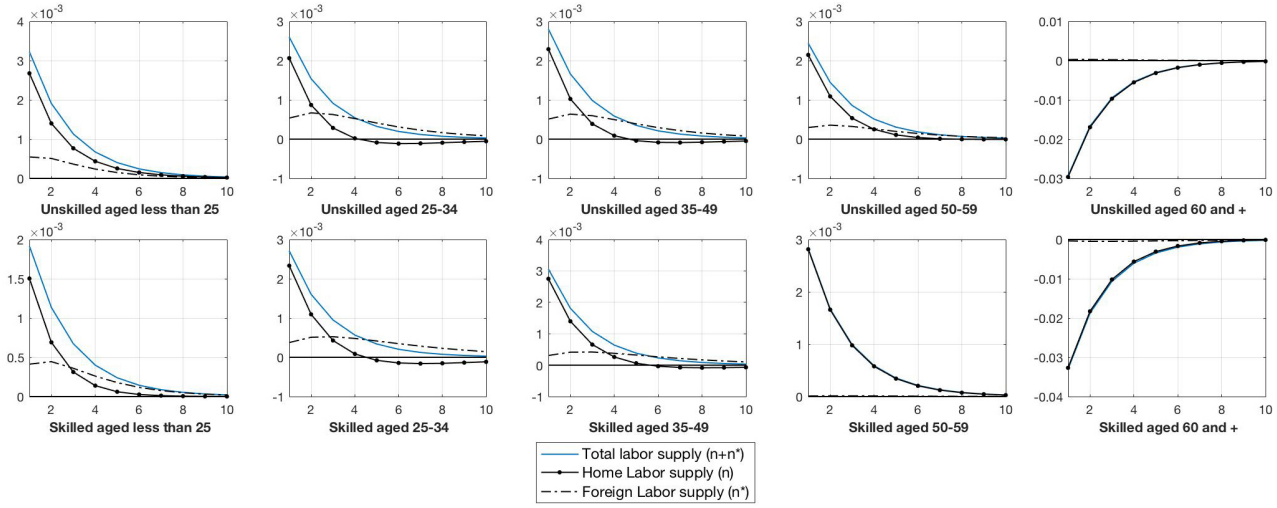
Next, the effects of the fiscal reform conducted in Uruguay after 2006 are analyzed. As mentioned above, Uruguay changed its structure of income taxation in 2007 from a flat tax rate on labor income to a progressive tax system. This reform, on average, decreased the marginal tax rate paid by unskilled workers and younger skilled workers and increased it for people over 50 years of age. At the same time, changes to the health care system were applied, changing the contributions to the NHF. Finally, the VAT tax rate was reduced from 23% to 22%, but improvements in audits of different kinds of taxes have increased the marginal effective VAT revenue, and thus the implicit tax rate paid by consumers.[28] These three changes are simulated as one tax reform in the calibrated benchmark model to

[28]The *Dirección General Impositiva* of Uruguay estimated that the implicit VAT tax rate generated government revenues of 10.2% of the GDP in 2012. In this model, that implies that the implicit VAT tax rate increased from 0.10 to 0.11.



Notes: Impulse response of total labor supply, labor supply at home and foreign, of unskilled and skilled age-groups to a 1% positive shock in education transfers. Results are depicted as percentage point deviations from the steady-state values.

Figure 1.4: Shock in education transfers



Notes: Impulse response of total labor supply, labor supply at home and foreign, of unskilled and skilled age-groups to a 1% positive shock in retirement pensions transfers. Results are depicted as percentage point deviations from the steady-state values.

Figure 1.5: Shock in retirement pensions transfers

address the effects in macroeconomic variables, specifically in the participation rates (home and abroad), and to assess the dynamic behind these results.[29] The marginal labor tax rate paid after the reform (including the change in marginal taxes and the contribution to NHF) is estimated using the Encuesta Continua de Hogares from the Instituto Nacional de Estadística (ECH-INE) from Uruguay of 2012.

In the last decade, public social spending in education, social security, and pensions has increased in Uruguay. Public social spending in education has risen from almost 3% of the GDP in 2006 to 4.5% in 2012. Social security and pensions also grew due to the changes described in subsection 2.1, and they increased by 1% of the GDP between 2006 and 2012. Two simulated scenarios of the increase in public spending on education and pensions were conducted.[30] To finance pensions and higher educational expenditure the marginal tax rates on labor were increased, in a way that allowed for new revenue to fund higher social spending.

The results of the changes in the three scenarios are presented in Tables 1.10 and 1.11.[31] The fiscal reform reduced the implicit labor tax rate for almost all age-ability groups, increasing after-tax incomes and the labor supply of these groups. The age-ability groups with a higher implicit tax rate on labor tend to diminish their labor supply at home. Emigration decreases in those groups who face a smaller income tax rate.[32] The second column of Table 1.11 reports changes of the main variables when the fiscal reform is applied. The after-tax wages increase, making staying at home more attractive; total employment at home rises, and total employment abroad decreases; in turn, total consumption at home increases due to the increase in total labor income at home. As the calibrated remittances are low, they do not play a significant role in the determination of total consumption at home. On the other hand, total consumption abroad is reduced, as fewer workers prefer to stay abroad.

The rise in educational spending and pensions, funded through distortionary taxes, have the same qualitative effect in both simulations. To finance a 1.5% (1%) increase in GDP in educational spending (pensions), an increase of 17.2% (11.5%) in the marginal tax rate is required. The higher taxes reduce the after-tax income, encouraging emigration and reducing labor supply at home. In the educational spending scenario, the reduction in the

[29]This scenario is referred to as "Fiscal Reform".

[30]The scenario of increasing social spending in education is referred to as "Education" and the scenario of the increase in social security and pensions as "Pensions".

[31]Table 1.A.2 in the Appendix reports the estimated marginal tax rates of the three analyzed scenarios.

[32]On average, the fiscal reform involved a reduction of 2% of the implied tax rate on labor incomes.

labor supply of young people is caused by a combination of a substitution effect (due to higher transfers) and by higher labor tax rates. Similarly, in the pensions scenario, the decline in the labor supply of elders is due to a combination of higher transfers and taxes. Total consumption at home falls because of reduced incomes and low remittances. On the contrary, consumption abroad increases as more members of the family decide to emigrate.

Table 1.10: Changes in simulated scenarios on participation rates from steady-state values

			Change from SS					
	Benchmark[a]		Fiscal Reform[b]		Education[c]		Pensions[d]	
	n_{ss}	n^*_{ss}	Δn_{FR}	Δn^*_{FR}	Δn_{Educ}	Δn^*_{Educ}	Δn_{Pen}	Δn^*_{Pen}
Unskilled								
0 to 24 years old	0.419	0.055	0.002	-0.002	-0.026	0.008	-0.001	0.005
25 to 34 years old	0.626	0.156	0.007	-0.008	-0.013	0.023	-0.008	0.014
35 to 49 years old	0.681	0.163	0.008	-0.009	-0.015	0.025	-0.009	0.015
50 to 59 years old	0.651	0.084	0.003	-0.004	-0.002	0.012	-0.002	0.007
60 years and older	0.185	0.045	0.002	-0.002	-0.001	0.006	-0.006	0.003
Skilled								
0 to 24 years old	0.348	0.071	0.005	-0.005	-0.016	0.012	-0.005	0.007
25 to 34 years old	0.565	0.251	0.012	-0.013	-0.042	0.049	-0.027	0.030
35 to 49 years old	0.642	0.279	0.001	-0.006	-0.061	0.067	-0.039	0.042
50 to 59 years old	0.553	0.298	-0.014	0.008	-0.076	0.082	-0.049	0.053
60 years and older	0.191	0.140	-0.014	0.012	-0.032	0.035	-0.028	0.024

Notes:

[a] Participation rates at home and abroad in the steady-state equilibrium.

[b] Changes from the steady-state equilibrium after the fiscal reform implemented in 2007 ($\Delta n_{FR} = n_{\text{fiscal reform}} - n_{\text{ss}}$ and $\Delta n^*_{FR} = n^*_{\text{fiscal reform}} - n^*_{\text{ss}}$).

[c] Changes from the steady-state equilibrium after a rise in education transfers of 1.5% of GDP ($\Delta n_{Educ} = n_{\text{education}} - n_{\text{ss}}$, and $\Delta n^*_{Educ} = n^*_{\text{education}} - n^*_{\text{ss}}$).

[d] Changes from the steady-state equilibrium after a rise in pension transfers of 1% of GDP ($\Delta n_{Pen} = n_{\text{pensions}} - n_{\text{ss}}$, and $\Delta n^*_{Pen} = n^*_{\text{pensions}} - n^*_{\text{ss}}$).

Table 1.11: Benchmark and simulated scenarios on the main variables of the model.

	Benchmark	Fiscal reform	Education	Pensions
Output	1.162	1.163	1.115	1.134
Wage at home	1.753	1.752	1.790	1.775
After-tax wage home	1.564	1.568	1.576	1.568
Total employment in home	0.470	0.473	0.448	0.460
Total employment in foreign	0.117	0.113	0.139	0.130
Total consumption in home	1.167	1.168	1.121	1.140
Total consumption in foreign	0.433	0.421	0.518	0.485
Stylized facts				
Pre-tax log wage difference (in %)	82.5%	82.6%	80.4%	81.3%
Remittances-GDP ratio	0.5%	0.4%	0.6%	0.5%
Government transfers to young-GDP ratio (in %)	2.9%	2.9%	4.5%	3.0%
Government transfers to elders-GDP ratio (in %)	7.0%	7.0%	7.3%	8.0%

Finally, it is worth mentioning that, because of the tax structure, the response to changes in taxes and transfers differ depending on age and ability. The skilled members of the household, who face higher implicit tax rates, are more likely to emigrate as their marginal tax rates increase. Consequently, redistribution of labor income taxes may encourage these ability groups to emigrate to avoid the tax burden.

1.6 Conclusion

The model presented in this paper tries to fill an existing gap in the literature by addressing the relationship between international migration and redistributive policies in the country of origin. The model also discusses the fact that migration and redistributive policies do not have the same impact on all age-ability groups. Furthermore, it explains the business cycle dynamics for aggregate variables as well as for the different groups. Migration decisions are the result of the optimization problem solved by the families, considering the wage differential and adjustment costs of migrating. Remittances are endogenously determined when individuals choose migration and consumption abroad. Exogenous government transfers to young people and the elderly allow an examination of the macroeconomic impact of changes in the transfer system. The model is calibrated using data on migration flows of

Uruguayans, government transfers to education, social security and pensions, and other macroeconomic variables. The estimated model matches the cyclical dynamic of targeted and untargeted moments correctly.

Based on this model, the results suggest that a tax reform that applies progressive tax rates, if the marginal change in the tax of high-income members of the economy is small enough, could generate an income effect that reduces emigration. Conversely, an increase in transfers to young people and the elderly financed with distortionary taxes produces a reduction in domestic wages, promoting emigration. Due to the tax and transfer structure of the economy, the migration process is not random among members of the household. Skilled agents in working ages are the ones that face higher migration rates. In terms of the life cycle, and taking into account the intergenerational transfer system, allocation of labor supply at home and abroad can be summarized as follows: when family members are young, they tend to choose low migration rates and allocate their time for studying taking full advantage of the transfer they receive. When they grow up and start their working life, some of the family members opt to migrate to avoid higher taxes that finance non-active members. At a retirement age, labor supply and migration decline due to the transfer the elderly receive in this stage of life.

Thus, the generous social policies for the young and the elderly enhance incentives to study and retire at home but work abroad. Increasing taxes on the active population that stays in Uruguay to finance these policies further incentivizes migration.

Bibliography

Aggarwal, R., DemirgÂ¸c-Kunt, A., and Martınez, M. (2011). "Do Remittances Promote Financial Development?". *Journal of Development Economics*, 96(2):255–264.

Alessandria, G., Bai, Y., and Deng, M. (2020). "Migration and Sovereign Default Risk". *Journal of Monetary Economics*, 113:1–22.

Annicchiarico, B. and Cesaroni, C. (2018). "Tax Reforms and the Underground Economy: A Simulation-Based Analysis". *International Tax and Public Finance*, 25(2):458–518.

Arseneau, D. and Chugh, S. (2012). "Tax Smoothing in Frictional Labor Markets". *Journal of Political Economy*, 120(5):926–985.

Bahadir, B., Chatterjee, S., and Lebesmuehlbacher, T. (2018). "The Macroeconomic Consequences of Remittances". *Journal of International Economics*, 111:214–232.

Bandeira, G., Caballé, J., and Vella, E. (2018). "Should I Stay or Should I Go? Austerity, Unemployment and Migration". *Working Paper No. 1839*, Banco de España.

Bandeira, G., Caballé, J., Vella, E., et al. (2019). "Fiscal Austerity and Greek Emigration: A missing link". Working Paper 1075, Barcelona GSE.

Barker, E. R. (2020). "The Resource Curse and Migration". Technical report, The Univeristy of Sheffield.

Batu, M. (2017). "International Worker Remittances and Economic Growth in a Real Business Cycle Framework". *Structural Change and Economic Dynamics*, 40:81 – 91.

Berry, R. and Soligo, R. (1969). "Some Welfare Aspects of International Migration. *Journal of Political Economy*, 77(5):778–794.

Borjas, G. (1999). "Immigration and Welfare Magnets". *Journal of Labor Economics*, 17(4):607–637.

Bucheli, M., Forteza, A., and Rossi, I. (2010). "Work Histories and the Access to Contributory Pensions: the Case of Uruguay". *Journal of Pension Economics & Finance*, 9(3):369–391.

Bucheli, M. and González, C. (2011). "Public Transfer Flows between Generations in Uruguay". In Lee, R. and Mason, A., editors, *Population Aging and the Generational Economy: A Global Perspective*. Edward Elgar Publishing.

Busato, F. and Chiarini, B. (2004). "Market and Underground Activities in a two-Sector Dynamic Equilibrium Model". *Economic Theory*, 23(4):831–861.

Busato, F., Chiarini, B., and Rey, G. M. (2012). "Equilibrium Implications of Fiscal Policy with Tax Evasion: A Long Run Perspective". *International Review of Law and Economics*, 32(2):197–214.

Cabella, W. and Pellegrino, A. (2005). "Una Estimación de la Emigración Internacional Uruguaya entre 1963 y 2004". *Working Paper, 70*, Facultad de Ciencias Sociales, Universidad de la República.

Casacuberta, C. and Bucheli, M. (2010). "Asistencia a Instituciones Educativas y Actividad Laboral de los Adolescentes en Uruguay, 1986-2008. *La Desafiliación en la Educación Media y Superior del Uruguay: conceptos, estudios y políticas*, Facultad de Ciencias Sociales, Universidad de la República.

CEPAL (2016). *"Panorama Social de América Latina, 2015"*. Comisión Económica para América Latina y el Caribe, (LC/G.2691-P), Santiago, 2016.

Chatterjee, S. and Turnovsky, S. (2018). "Remittances and the Informal Economy". *Journal of Development Economics*, 133:66–83.

Chiquiar, D. and Hanson, G. (2005). "International Migration, Self-Selection, and the Distribution of Wages: Evidence from Mexico and the United States". *Journal of Political Economy*, 113(2):239–281.

Cid, A. and Ferrés, D. (2010). "The Coexistence of Degree-premium and High-Dropout Rates in the Uruguayan Secondary Education: An Incentives Problem". *Social Science Research Network*, SSRN 1850894.

de Armas, G. (2018). "Trayectorias Educativas en Uruguay. Principales Rasgos, Tendencias y Desafíos para las Políticas Públicas". UNICEF.

Djajić, S. (1986). "International Migration, Remittances and Welfare in a Dependent Economy". *Journal of Development Economics*, 21(2):229–234.

Forteza, A., Apella, I., Fajnzylber, E., Grushka, C., Rossi, I., and Sanroman, G. (2009). "Work Histories and Pension Entitlements in Argentina, Chile and Uruguay". *World Bank Policy Research Working Paper*, No. 0926.

Fromentin, V. (2017). "The Long-Run and Short-Run Impacts of Remittances on Financial Development in Developing Countries". *The Quarterly Review of Economics and Finance*, 66:192–201.

Garcia-Milà, T., Marcet, A., and Ventura, E. (2010). "Supply Side Interventions and Redistribution". *The Economic Journal*, 120(543):105–130.

Giulietti, C. (2014). "The Welfare Magnet Hypothesis and the Welfare Take-up of Migrants". *IZA World of Labor*.

Hanson, G. (2010). "International Migration and the Developing World". In Rodrick, D. and Rosenzweig, M., editors, *Handbook of development economics*, volume 5, pages 4363–4414. Elsevier.

Iqbal, K. and Turnovsky, S. J. (2008). Intergenerational Allocation of Government Expenditures: Externalities and Optimal Taxation. *Journal of Public Economic Theory*, 10(1):27–53.

Jakubiak, I. (2017). "Migration and Welfare Systems-State of the Art and Research Challenges". *Central European Economic Journal*, 1(48):51–70.

Janiak, A. and Monteiro, P. (2010). "Labor Force Heterogeneity: Implications for the Relation Between Aggregate Volatility and Government Size". *Working Paper 272*, Universidad de Chile.

Katz, G., Barrenechea, P., and Pastori, H. (2004). "Construcción de una Matriz de Contabilidad Social para Uruguay para el año 2000". *Working Paper, 20/04*, Facultad de Ciencias Sociales, Universidad de la República.

Koolhaas, M. and Pellegrino, A. (2009).

Kydland, F. (1984). "Labor-Force Heterogeneity and the Business Cycle". In *Carnegie-Rochester Conference Series on Public Policy*, volume 21, pages 173–208. Elsevier.

Lavalleja, M., Rossi, I., and Tenenbaum, V. (2018). "La Flexibilización de las Condiciones de Acceso a las Jubilaciones en el Uruguay". *Estudios y Perspectivas*, 33, CEPAL.

Lavalleja, M. and Tenenbaum, V. (2017). "Seguridad Social en el Uruguay: un Análisis de los Resultados de la Ley de Flexibilización del Acceso a las Jubilaciones". *Estudios y Perspectivas*, 30, CEPAL.

Lazaretou, S. (2016). "The Greek Brain drain: the new pattern of Greek emigration during the recent crisis". *Economic Bulletin*, (43):31–53.

Lee, R. and Mason, A. (2011). *"Population Aging and the Generational Economy: A Global Perspective"*. Edward Elgar Publishing.

Leers, T., Meijdam, L., and Verbon, H. (2004). "Ageing, Migration and Endogenous Public Pensions". *Journal of Public Economics*, 88(1-2):131–159.

Lim, S. and Morshed, M. (2017). "Fiscal Policy in a Small Open Economy with Cross-Border Labor Mobility". *Journal of Macroeconomics*, 52:147–174.

Lozej, M. (2019). "Economic Migration and Business Cycles in a Small Open Economy with Matching Frictions". *Economic Modelling*, 81:604–620.

Maliar, L. and Maliar, S. (2000). "Differential Responses of Labor Supply Across Productivity Groups". *Journal of Macroeconomics*, 22(1):85–108.

Maliar, L. and Maliar, S. (2001). "Heterogeneity in Capital and Skills in a Neoclassical Stochastic Growth Model". *Journal of Economic Dynamics and Control*, 25(9):1367–1397.

Maliar, L. and Maliar, S. (2003). "The Representative Consumer in the Neoclassical Growth Model with Idiosyncratic Shocks". *Review of Economic Dynamics*, 6(2):362–380.

Mandelman, F. and Zlate, A. (2012). "Immigration, Remittances and Business Cycles". *Journal of Monetary Economics*, 59(2):196–213.

Moffitt, R. (1992). "Incentive Effects of the US Welfare System: A Review". *Journal of Economic Literature*, 30(1):1–61.

Mountford, A. (1997). "Can a Brain Drain Be Good for Growth in the Source Economy?". *Journal of Development Economics*, 53(2):287–303.

OECD (2017). *"Revenue Statistics in Latin America: Comparative tables"*. OECD Tax Statistics (database).

Orsi, R., Raggi, D., and Turino, F. (2014). "Size, Trend, and Policy Implications of the Underground Economy". *Review of Economic Dynamics*, 17(3):417–436.

Pellegrino, A. (2009). "Uruguay: País de Migrantes Internos y Externos". *Material de divulgación*, UNDP.

Pellegrino, A. and Koolhaas, M. (2008). "Migración Internacional: Los Hogares de los Emigrantes". In Varela, C., editor, *"Demografıa de una Sociedad en Transición: La Población Uruguaya a Inicios del Siglo XXI"*. Chapter 5, United Nations Population Fund.

Pellegrino, A. and Vigorito, A. (2009). La emigración calificada desde América Latina y las iniciativas nacionales de vinculación. Un análisis del caso Uruguayo. *Pensamiento Iberoamericano*, (4):189–215.

Peterman, W. (2016). "Reconciling Micro and Macro Estimates of the Frisch Labor Supply Elasticity". *Economic Inquiry*, 54(1):100–120.

Quibria, M. (1997). "International Migration, Remittances and Income Distribution in the Source Country: A Synthesis". *Bulletin of Economic Research*, 49(1):29–46.

Sand, E. and Razin, A. (2006). "Immigration and the Survival of Social Security: A Political Economy Model". Working Paper 12800, National Bureau of Economic Research.

Sand, E. and Razin, A. (2007). "The Political-Economy Positive Role of the Social Security System in Sustaining Immigration (But Not Vice Versa)". Working Paper 13598, National Bureau of Economic Research.

Schmitt-Grohé, S. and Uribe, M. (2004). "Solving Dynamic General Equilibrium Models Using a Second-Order Approximation to the Policy Function". *Journal of Economic Dynamics and Control*, 28(4):755–775.

Stark, O., Helmenstein, C., and Yegorov, Y. (1997). "Migrants' Savings, Purchasing Power Parity, and the Optimal Duration of Migration". *International Tax and Public Finance*, 4(3):307–324.

Stark, O. and Taylor, J. (1991). "Migration Incentives, Migration Types: The Role of Relative Deprivation". *The Economic Journal*, 101(408):1163–1178.

Storesletten, K. (2000). "Sustaining Fiscal Policy through Immigration". *Journal of Political Economy*, 108(2):300–323.

United Nations (2017). "International Migration Report 2017". United Nations.

Végh, C., Lederman, D., and Bennett, F. (2017). "Leaning Against the Wind: Fiscal Policy in Latin America and the Caribbean". The World Bank.

1.A Appendix for Chapter 1

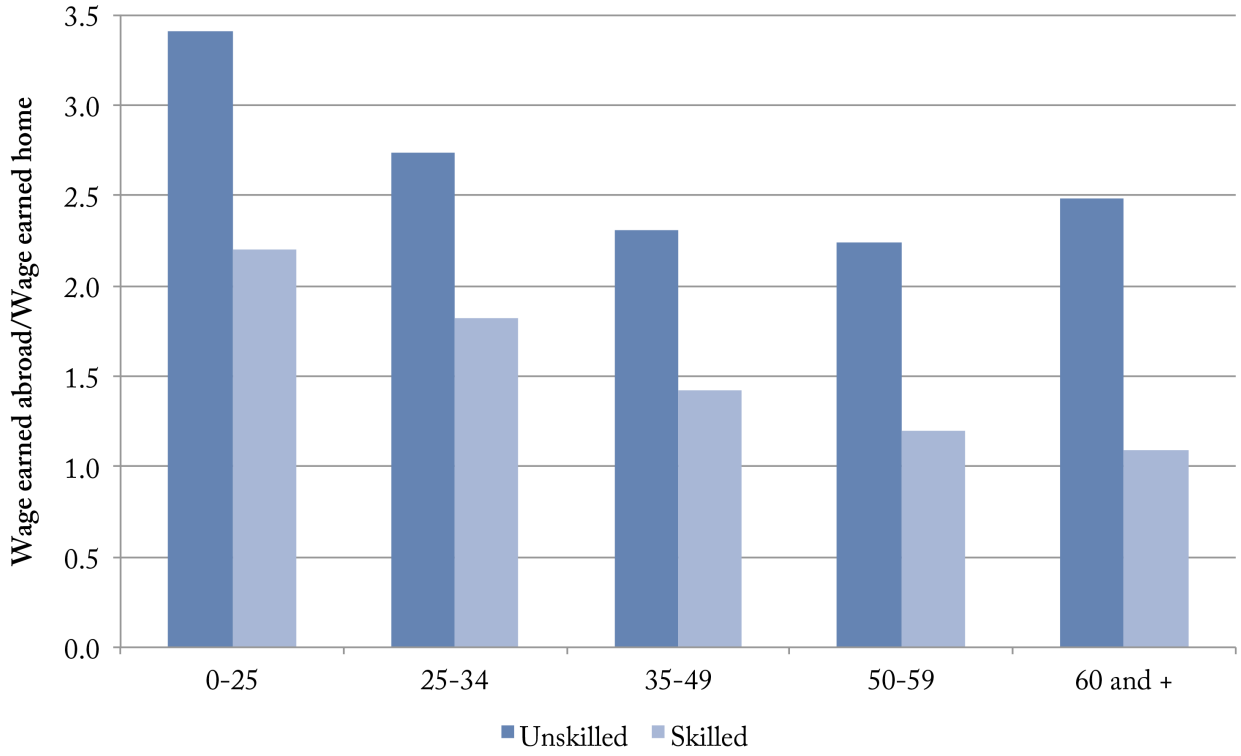


Figure 1.A.1: Mean pre-tax wage earned from Uruguayans living abroad, relative to those Uruguayans living in Uruguay (in 2006 PPP values)

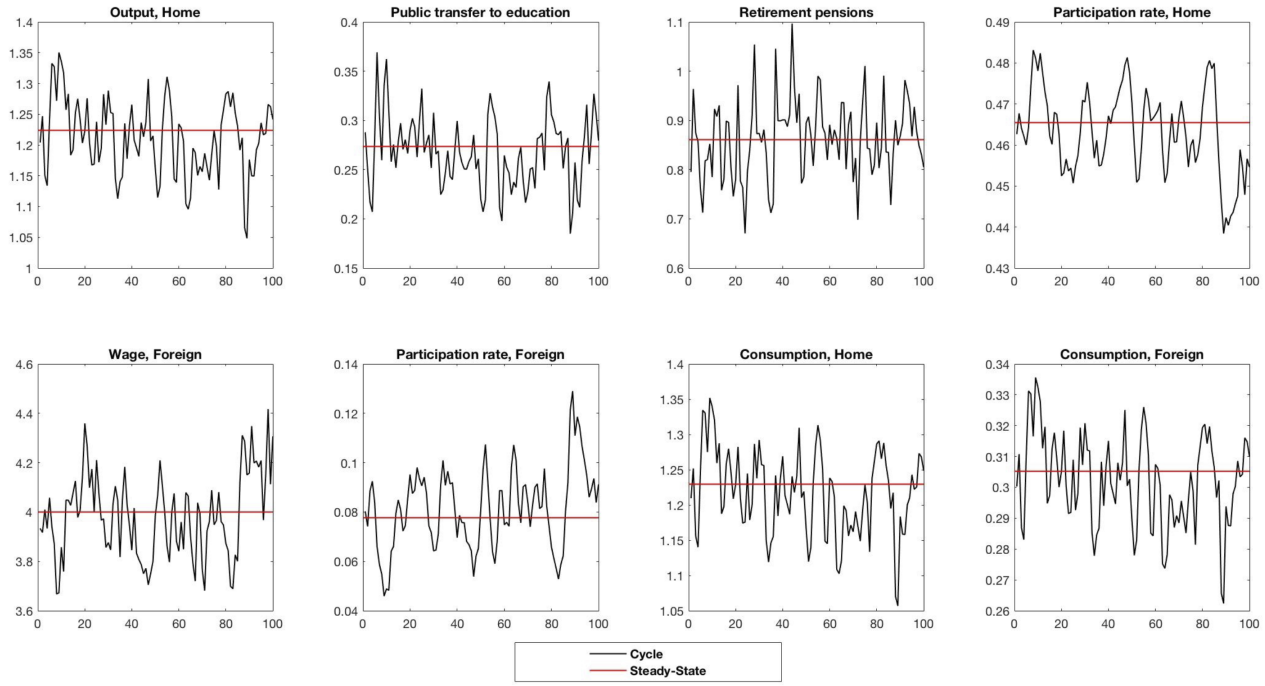


Figure 1.A.2: Simulation of main variables of the model, cycle and steady-state values

Table 1.A.1: Correlation between GDP and participation rates at home and foreign.

	Participation rate at home		Participation rate at foreign	
	Data	Benchmark	Data	Benchmark
Unskilled				
0 to 24 years old	0.769	-0.784	0.177	-0.817
25 to 34 years old	0.628	0.818	-0.298	-0.759
35 to 49 years old	0.327	0.824	-0.528	-0.765
50 to 59 years old	0.353	0.887	-0.566	-0.781
60 years and older	0.461	0.120	-0.064	-0.812
Skilled				
0 to 24 years old	0.466	0.747	0.259	-0.808
25 to 34 years old	0.673	0.739	-0.203	-0.710
35 to 49 years old	0.570	0.767	-0.380	-0.737
50 to 59 years old	0.526	0.765	-0.356	-0.742
60 years and older	0.261	0.784	-0.062	-0.778

Table 1.A.2: Marginal tax rates applied in the benchmark model and simulated scenarios.

	Benchmark	Fiscal reform	Education	Pensions
Unskilled				
0 to 24 years old	0.096	0.090	0.113	0.107
25 to 34 years old	0.117	0.110	0.138	0.131
35 to 49 years old	0.123	0.115	0.144	0.137
50 to 59 years old	0.105	0.098	0.123	0.117
60 years and older	0.068	0.064	0.079	0.075
Skilled				
0 to 24 years old	0.148	0.136	0.174	0.165
25 to 34 years old	0.176	0.170	0.206	0.196
35 to 49 years old	0.192	0.195	0.225	0.214
50 to 59 years old	0.178	0.185	0.208	0.198
60 years and older	0.140	0.148	0.164	0.156

Chapter 2

Middle-term Effects of Sibling's Sex Composition on Early Childhood Development

2.1 Introduction

One of the most important social interactions in people's lives is the one with siblings. This peer-to-peer influence is a result of the direct effect between them, and the indirect effect caused by parental behavior. Research on children's sex composition shows that parental behavior towards boys and girls is different but does not go further into analyzing if this different treatment from parents persists when they have same-sex or opposite-sex children. If there are differences between the treatment of boys and girls, the sex of his or her siblings may reinforce or offset the indirect effects. McHale et al. (2003) have suggested that within-family comparison of how parents treat sisters versus brothers is a powerful test to detect the real behavior of parents towards boys and girls.

This study aims to analyze the middle-term effects of siblings' sex composition on childhood development and to isolate possible channels that may explain these results. Cognitive and non-cognitive development in early childhood plays a central role in an individual's successful life. Cunha and Heckman (2007) explain skill formation as a dynamic process, where investments made by parents, even before the child is born, are complements, and there is investment self-productivity. They also show that later investments in childhood compensate for some of the early disadvantages, although they are less effective and costly. Cognitive skills concern problem-solving and communication abilities, whereas

non-cognitive skills range from time preference, risk aversion, determination, motivation, emotional self-control, to stable self-esteem. Both aspects of development are essential predictors of schooling attainment, human capital formation, labor market outcomes, health, marriage, and divorce (Heckman and Rubinstein, 2001; Carneiro et al., 2007; Conti and Heckman, 2012; Fiorini and Keane, 2014).

Social scientists have long been interested in how siblings affect each other and how they affect parental behavior. The peer effect is related to the direct role siblings play as the primary socialization environment. If girls perform better (worse) in some areas than boys, same-sex siblings will boost (disrupt) their development. The parental behavior effect can be linked to the preferences of parents over boys or girls, or to constraints due to economies of scale in time and money from having a same-sex or mixed-sex composition of their children. The main effect of parental preferences regarding their children's sex composition is over fertility choices. When parents have a preference for mixed-sex composition, the sex of the second-born child may affect the decision to have another child (Angrist and Evans, 1998). This fertility channel affects the family size in which children grow up, impacting on the constraints that parents face. The model developed by Becker and Lewis (1973), have indicated that parents derive utility from the quantity and quality of their children, but they face a "quantity-quality trade-off": due to the limited resources in the household, a higher number of siblings in the family leaves each child with fewer resources of time and money.[1]

The main contribution of this paper to the literature is to explain the differences observed between childhood development of first-born boys and first-born girls, who have a second-born sibling of the same-sex or opposite-sex. The results also contribute to the growing literature on siblings' sex composition by analyzing the mechanisms leading to these results, and to the child-gender literature by showing that parents behave differently regarding their sons and daughters.

The data used in this paper is from the *"Encuesta de Nutrición, Desarrollo Infantil y Salud"* (ENDIS) of Uruguay, a panel survey conducted on a representative sample of children under six years old in Uruguay. This survey has information that allows the identification of the causal effect of siblings' sex composition and includes information on motor skills, cognitive and non-cognitive development, as well as information on family structure and socioeconomic characteristics of the household.

The central hypothesis is that the younger sibling has a different impact on the older sib-

[1]The "resource dilution" hypothesis, Powell and Steelman (1993) propose a very similar mechanism of limited resources in the household.

ling's development, depending on his or her sex. The expected effect comes from different parental behavior on rearing boys and girls and changes in the constraints the parents face.

The empirical strategy used in this paper relies on the randomness of the sex of the second-born child in order to identify the causal effects of siblings' sex composition on motor skills, as well as cognitive and non-cognitive childhood development. The causal effect is estimated separately for boys and girls, using regression analysis. The data is restricted to first-born children with at least one younger sibling. The evidence shows that girls are not affected in the middle-term by siblings, whichever the sex. However, first-born boys with a younger brother have significant negative effects on motor skills and non-cognitive abilities, compared to those exposed to a younger sister.

This work studies several channels to explain these results: the family structure, parental investment on quality time, schooling and health, and parental practices. The evidence suggests that parental quality time investment is the most relevant variable to explain results. There are also significant differences between first-born boys with same and opposite-sex second-born siblings, in terms of the likelihood of living with both parents and attending school.

The outline of the paper is as follows: Section 2.2 reviews prior research on this topic and describes possible channels to explain results, Section 2.3 describes the data, shows some descriptive statistics and outlines the identification strategy, Section 2.4 presents the main results on childhood development, and analyzes potential mechanisms of transmission and Section 2.5 concludes.

2.2 Related Literature

The next two subsections present a brief overview of the most relevant works on siblings' sex composition and on the potential mechanisms that can be drivers of such differences. First, the evidence on siblings' sex composition on long term outcomes is summarized, as well as the evidence on childhood development. Secondly, the expected effects on child development related to parents' behavior are discussed, and a summary of the evidence regarding parents' behavior depending on the sex composition of their children is made; including fertility decisions, family formation, labor-market activities of parents, parental investments and expectations, and parental practices.

2.2.1 Sex composition effect

The role of family size, birth order, child spacing, and siblings' sex composition on childhood development has been studied mainly in psychology, sociology, and economics. Recent studies focus on the causal effect of siblings' sex composition on long-term outcomes such as human capital accumulation, labor-market outcomes, occupational and partner choice, and family formation. Brenøe (2018) analyzed the cases of Danish first-born women with a younger sibling in the period from 1980 to 2016, and found that first-born women with a second-born brother have earnings that are 2% lower than those of women with a second-born sister; whereas Cools and Patacchini (2019), using data of women from the United States born on the late 1970s to early 1980s, detected that this reduction reaches 7%. Rao and Chatterjee (2018), using data from the National Longitudinal Survey of Youth of 1979 from the United States, show that first-born men who have a second-born brother have higher earnings than first-born men with a second-born sister, but educational attainment is not affected by their siblings' sex composition. Similarly, Peter et al. (2018), using Swedish data on siblings born between 1938 and 1970, found that first-born men who have a same-sex younger sibling increase their earnings, are more likely to get married, and have any children than men with an opposite-sex second-born siblings.[2] Vladasel (2018), using the same Swedish register data, found direct effects of siblings' sex composition on entrepreneurship when the father is present in the household. Other studies have analyzed the impact of being raised with more brothers or sisters on the educational attainment of men and women, and have shown that women reared with brothers reached, on average, more years of schooling than those with any sister (Butcher and Case, 1994).

The present study relates to Cyron et al. (2017), who analyzed the causal effect of having a younger sister on cognitive and non-cognitive development among six year old children in the United States. They found that having a younger sister (compared to having a younger brother) improves the results of boys but does not have any significant effects on girls. These authors do not analyze the channels behind their results, which is the central focus of this paper.

2.2.2 Potential mechanisms

Parents may have preferences over the number of children they wish to have as well as their sex composition. Conditional on having a child, the sex of the first-born could lead

[2]Results for women are lower and less robust than those indicated for men.

to parents having another child if the sex of the first-born is not in line with their preferences, changing fertility choices and the total number of children they want to have. This has a direct effect on parents' resources of time and money to rear their children. This idea was formalized by Becker and Lewis (1973) and Becker and Tomes (1976), who developed the "quantity-quality trade-off" model, predicting that the family size would have negative effects on different components of child development, as well as long term outputs.[3] Nonetheless, at present there is no robust evidence of a "quantity-quality trade-off" operating on middle or long-term outcomes (Black et al., 2005; Angrist et al., 2010; Black et al., 2010; Åslund and Grönqvist, 2010; Juhn et al., 2015; Guo et al., 2017).

Empirical evidence has been used to analyze the "first-born sex effect" and the "siblings' sex composition effect" on fertility choices. The literature found that parents have son preferences, meaning that the probability of having another child after a first-born male is lower than a first-born female (Angrist and Evans, 1998; Rosenzweig and Wolpin, 2000; Lundberg, 2005; Dahl and Moretti, 2008; Nguyen, 2019; Cools and Patacchini, 2019; Peter et al., 2018). Then, first-born males and first-two born siblings of mixed sexes are more likely to live in smaller households, which has a direct effect on parents' resources of time and money, required to rear their children.

The sex composition of the children can dilute the quantity-quality trade-off through economies of scale, for example, when families with same-sex children can benefit from buying fewer gender-specific goods. More resources allow for better development in children, which implies that in the medium term the effects of having a sibling of the same sex should be positive. There is evidence that this channel is important for India but not for The United States or Sweden (Rosenzweig and Wolpin, 2000; Angrist et al., 2010; Peter et al., 2018).

The impact of family structure and home environment on the future outcomes of children has already been recognized by Cunha and Heckman (2007), Heckman (2008), Currie and Almond (2011) and Heckman et al. (2013), among others. Family structure refers to the people who raise the children, and who directly affect the inputs of children's production function: the amount of money that parents are able to invest, the time they can spend with their children and the environment in which they grow up. Recent research on these top-

[3] Becker and Tomes (1976) have suggested that parents have to choose how to the investment in their children at a very young age, before knowing the skills they will need to enter the job market. Parents, in order to maximize their lifetime utility, are willing to invest more in the children who have a higher return rate. If the gender pay gap in the labor market favors boys, parents will invest more in their sons rather than in their daughters. This approach looks at the results of each gender separately, but these decisions may change depending on the sex composition of their children.

ics found that the absence of the father has negative effects on non-cognitive development, this effect being stronger if the absence occurs during early childhood, and boys are more affected by this absence than girls (McLanahan et al., 2013; Bertrand and Pan, 2013).

Literature related to the "first-born sex effect" on family-structure shows that parents with a first-born girl are more likely to get divorced than parents who have a first-born boy (Dahl and Moretti, 2008; Depew and Price, 2018; Reynolds, 2018). Saarela and Finnäs (2014) and Reynolds (2018) have analyzed the sex composition effect on family structure, but focusing on the differences between families with sons only and families with daughters only. These studies found that parents of only boys are more likely to be part of a union than parents of only girls. Cools and Patacchini (2019) do not find significant differences between girls with a same-sex or opposite-sex sibling on the probability of living with two biological parents.[4]

Parental time investment and parents' expectations on child development are significant issues that can lead to an improvement in academic achievement in developed countries (Del Boca et al., 2013; Fiorini and Keane, 2014; Bono et al., 2016; Nollenberger et al., 2016; Rodríguez-Planas and Nollenberger, 2018).[5] Differences within the two sexes have been analyzed, showing that boys require more investment in parental quality time than girls to reach similar results on child development (Ginja et al., 2017; Bertrand and Pan, 2013; Autor et al., 2019; Fan et al., 2015; Brenøe and Lundberg, 2018). Regarding the sex composition effect, some studies have found that providing extra time for parents to spend with their children has a significant impact on the child development of first-born when second-born is a boy (Sayour, 2019). Other studies have shown that parents spend more quality time with their daughters or have higher educational aspirations for girls (Bertrand and Pan, 2013; Baker and Milligan, 2016; Autor et al., 2019; Reynolds and Burge, 2008; Fortin et al., 2015). These authors claim that these changes in parental practices can account for the reverse gender gap in drop-out rates, high-school or college attendance.

The sex composition of children can also affect parental time allocation. According to preference- and constrained- based models, mothers spend more time with their daughters and fathers with their sons (Lundberg, 2005; Bertrand and Pan, 2013; Baker and Milligan, 2016). Following these models, parents could reproduce some gender-typed activities, affecting the household's division of labor and working choices. The literature related to the "first-born boys effect" found that first-born boys increase the gender specialization

[4] Cools and Patacchini (2019) do not present results for boys on this aspect.

[5] Parental time investment refers to the activities parents carry out with their children. For example, reading books, telling stories, singing songs, teaching new games, or playing with them.

in households with low educational level, but reduces it in families with highly educated parents (Lundberg, 2005; Pollmann-Schult, 2017; Fan et al., 2018).

Parenting stands for every choice made by parents and all the interactions they have with their children. These choices range from the very basic provision of food and health to supporting and promoting their intellectual development (**?**). Three types of parenting styles are usually defined to analyze how differences in home environment can affect child development: authoritarian, authoritative and permissive.[6] Several psychology studies have analyzed sex differences in parenting over early childhood development and found that the cognitive and non-cognitive development of boys is more sensitive to the home environment than girls'. Additionally, boys with siblings perform worse than boys without them (Mileva-Seitz et al., 2015; Golding and Fitzgerald, 2019; Kim et al., 2018). Less authoritarian and more authoritative parenting styles are usually associated with higher levels of childhood development (**?**Perazzo et al., 2018). This dimension has not been studied considering the sex composition of children.

Each of the mechanisms discussed in this section will be further analyzed in Section 2.4.2 to disentangle the causes of the differences found.

2.3 Data and Identification Strategy

2.3.1 Data

The source data used is the *"Encuesta de Nutrición, Desarrollo Infantil y Salud"* (ENDIS) of Uruguay. The ENDIS is a panel-data survey with a wide range of information on child well-being such as nutritional status, health, and development. The first round of the ENDIS was surveyed between 2013 and 2014 and is representative of children between 0 and 4 years of age. The data contains information on 3.077 children and 2.665 adults (97% mothers). The second round was performed between 2015 and 2016 and contains information on 2.611 children and 2.325 adults (96% mothers). The panel-data can be combined with the *"Encuesta Continua de Hogares"* (ECH) of Uruguay of 2012/2013, which contains the full socio-economic environment of the household.

The data used in the analysis is restricted to first-born children with a younger singleton sibling and excludes step or adopted siblings. It includes 133 first-born girls and 158 first-

[6] Doepke and Zilibotti (2017) show that parenting styles -authoritarian, authoritative and permissive- emerge as outcome variables from the maximization problem of parents and that each one is affected by parental preferences and the socioeconomic environment in which they live.

born boys.[7] The main reason to restrict the data is that an estimate on first-born children is the most straightforward and unbiased estimate of sibling sex composition on child development. If all children with a younger sibling are considered, the birth-order effect can bias the result, because parents with probably different preferences over the sex composition of their children are being compared. For the same reason, the analysis is carried out separately for girls and boys. Peter et al. (2018) support that this strategy helps to overcome the selection bias problem of parental preferences over the sex composition of children. If this is the case, results concerning first-borns represent the upper bound of the siblings' sex composition effect, since these children are the only ones who receive the full attention and investment of their parents before they have another child. Other unobservable variables could bias the results on higher-order siblings.

In this paper, the outcome variables are the components of the Ages and Stages Questionnaires (ASQ-3) and the Child Behavior Checklist (CBCL) tests. The ASQ-3 test was performed on children between one and 66 months old, and the first round was only administered in Montevideo (the capital city). The ASQ-3 test is internationally used to monitor child development progress. It contains five components that approximate the level of motor skills, cognitive, and non-cognitive development. Motor skills are measured by gross motor skills (walk, jump or run) and fine motor skills (hand-eye coordination and hand dexterity). Cognitive skills are measured through problem-solving (count and logical issues) and communication skills (speaking and listening abilities). Last, the non-cognitive skills are approximated by the socio-personal skills (the capacity to interact with other children, as well as the capability to wash and dress him or herself). These tests were performed by a qualified interviewer who asked the child to do the assigned tasks.[8]

The CBCL is used to evaluate the non-cognitive development of children through emotional and behavioral problems. It consists of 100 questions, scored with zero if the problem is absent, with one if the problem occurs sometimes, and two if the problem occurs often. Through these questions, two empirical syndromes called internalized and externalized

[7] In addition, results for the complete sample of children with younger siblings are in Appendix 2.A (224 girls and 260 boys).

[8] The test incorporates developmental differences, based on the age (in months) of the children. Each questionnaire consists of thirty questions, six for each component, which gives tasks that the child may or may not achieve. If the child achieves the assigned task, it is valued with ten points. If the child reaches it partially, he or she is assigned five points or zero points if the child cannot perform the task. The final score adds up each item of each area and is then standardized according to the child's age. If there are no conditions to perform a particular task for each component, it is imputed as missing. Fine-motor skills require the child to be in certain conditions, otherwise it can be harder to measure. Hence, this component has a high number of missing values.

problems can be detected. The groups of syndromes that integrate the internalized problems include the child being emotionally reactive, anxiety, depression, somatic complaints, and withdrawal (36 items). Externalized problems are a combination of attention problems and aggressive behavior (25 items). Finally, the total problem scale is the sum of all the items in the test. The five components of ASQ-3 and the three of CBCL are standardized, using the whole sample in each round, to have mean zero and standard deviation of one. While the ASQ-3 test measures skills, the CBCL measures problems. Thus, a higher value of the ASQ-3 test means higher development, whereas a higher value of the CBCL means lower development.[9] To avoid confusion in the interpretation of the coefficients of the ASQ-3 and CBCL tests, the estimation of the CBCL test will be presented with an opposite value from the one in the standardized test.

The data was balanced through outcome variables, dropping all observations with no information on the communication, gross-motor, problem-solving, and socio-personal measures of the ASQ-3 test.[10]-[11]-[12]

The average score for first-born girls and boys on the main outcome variables used in the regression analysis are presented in Table 2.1.[13] The outcome variables are grouped to approximate motor skills with fine and gross motor skills, cognitive development with communication and problem-solving skills, and non-cognitive development with socio-personal skills, the externalized, internalized, and total problem measures. On average, girls outperform boys on all tests. The most significant differences are on fine-motor skills where girls reach an average score of 0.405 and boys -0.276, and on problem-solving skills, where the score of girls is 0.259 and boys' is -0.09. Significant differences are also found in socio-personal skills, 0.241 for girls and -0.131 for boys, and in externalized problems, 0.058 versus -0.170.

[9] Details of the standardization method are explained in Perazzo et al. (2018).

[10] Children who have a score of 0 on all ASQ-3 tests and 0 problems of the CBCL test has been removed from the sample.

[11] Fine-motor skills are not balanced across due to the small number of observations.

[12] The CBCL sub-sample could have been considered, since, in the second round, it gave information on 37 first-born girls and 26 first-born boys more than the ASQ-3 sample. Results of CBCL tests and mechanisms of this sub-sample are not statistically different from the results shown in this paper.

[13] Table 2.A.1 of the Appendix presents these summary statistics for the whole sample of children with a younger sibling.

Table 2.1: Summary statistics by sex.

	Girls	Boys	Difference
Motor skills			
Gross-motor skills	0.064	-0.030	-0.095
	(0.674)	(0.942)	(0.098)
Fine-motor skills	0.405	-0.276	-0.681
	(0.689)	(1.085)	(0.141)
Cognitive skills			
Communication	0.180	-0.017	-0.197
	(0.743)	(0.985)	(0.104)
Problem-solving	0.259	-0.090	-0.350
	(0.783)	(1.134)	(0.116)
Non-cognitive skills			
Socio-emotional	0.241	-0.131	-0.372
	(0.656)	(0.941)	(0.097)
Externalized problems	0.058	-0.170	-0.229
	(0.874)	(0.980)	(0.110)
Internalized problems	0.041	-0.093	-0.134
	(0.79)	(0.834)	(0.096)
Total problems	0.041	-0.107	-0.148
	(0.795)	(0.863)	(0.098)
Observations	133	158	

Notes: The table reports means and standard deviations of standardized test scores of the ASQ-3 and CBCL tests. The sample consists of first-born children who have a younger sibling in wave 2. The sample size of fine-motor skills for girls is 79 and for boys 95.

Figures 2.1 to 2.3 show the average motor skills, cognitive, and non-cognitive development scores for girls and boys who have a same-sex or an opposite-sex second-born sibling. The median score for girls is higher than that of boys, irrespective of the sex of the younger sibling. The differences between same-sex and opposite-sex siblings are more significant among boys than girls, and they also have more dispersion. The most critical differences between same-sex and opposite-sex siblings are on fine-motor skills for both sexes and the different measures of non-cognitive development among boys.

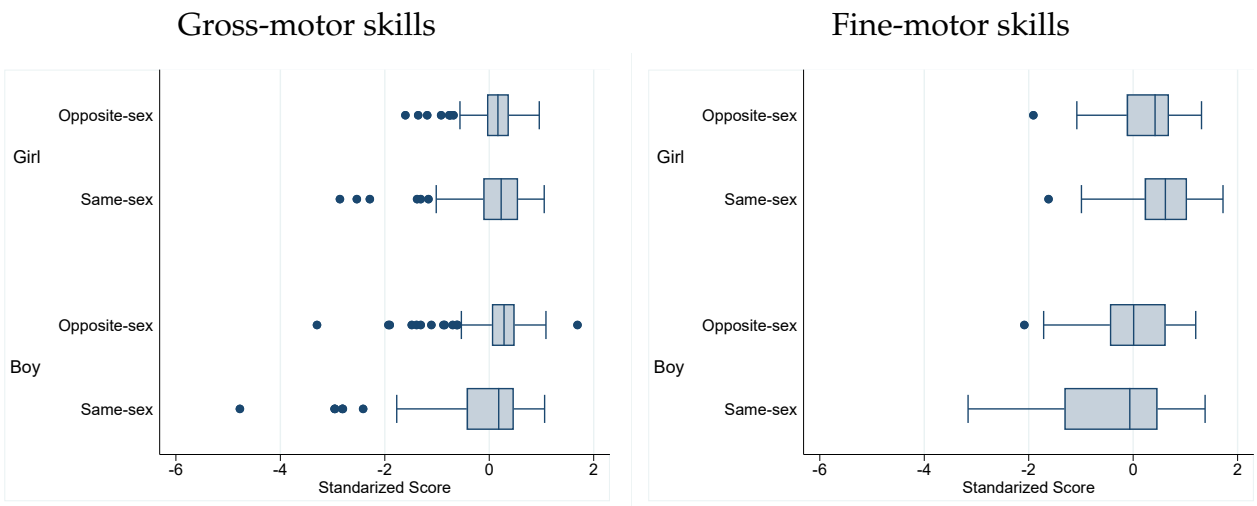


Figure 2.1: Motor Skills distributions for first-born children with an opposite-sex or a same-sex second-born sibling, by the sex of the first-born.

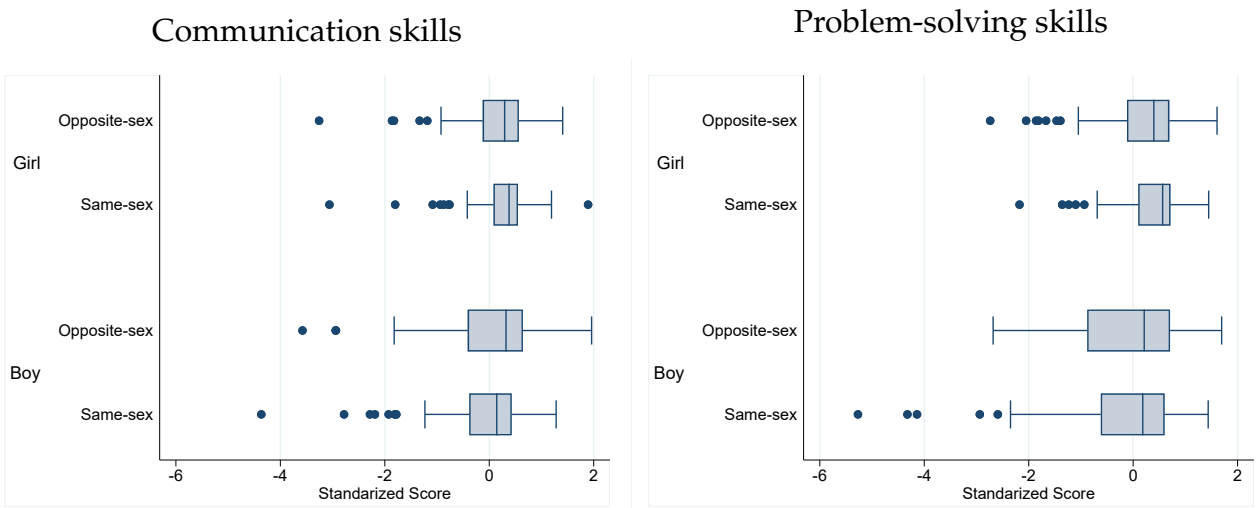


Figure 2.2: Cognitive Skills distributions for first-born children with an opposite-sex or a same-sex second-born sibling, by the sex of the first-born.

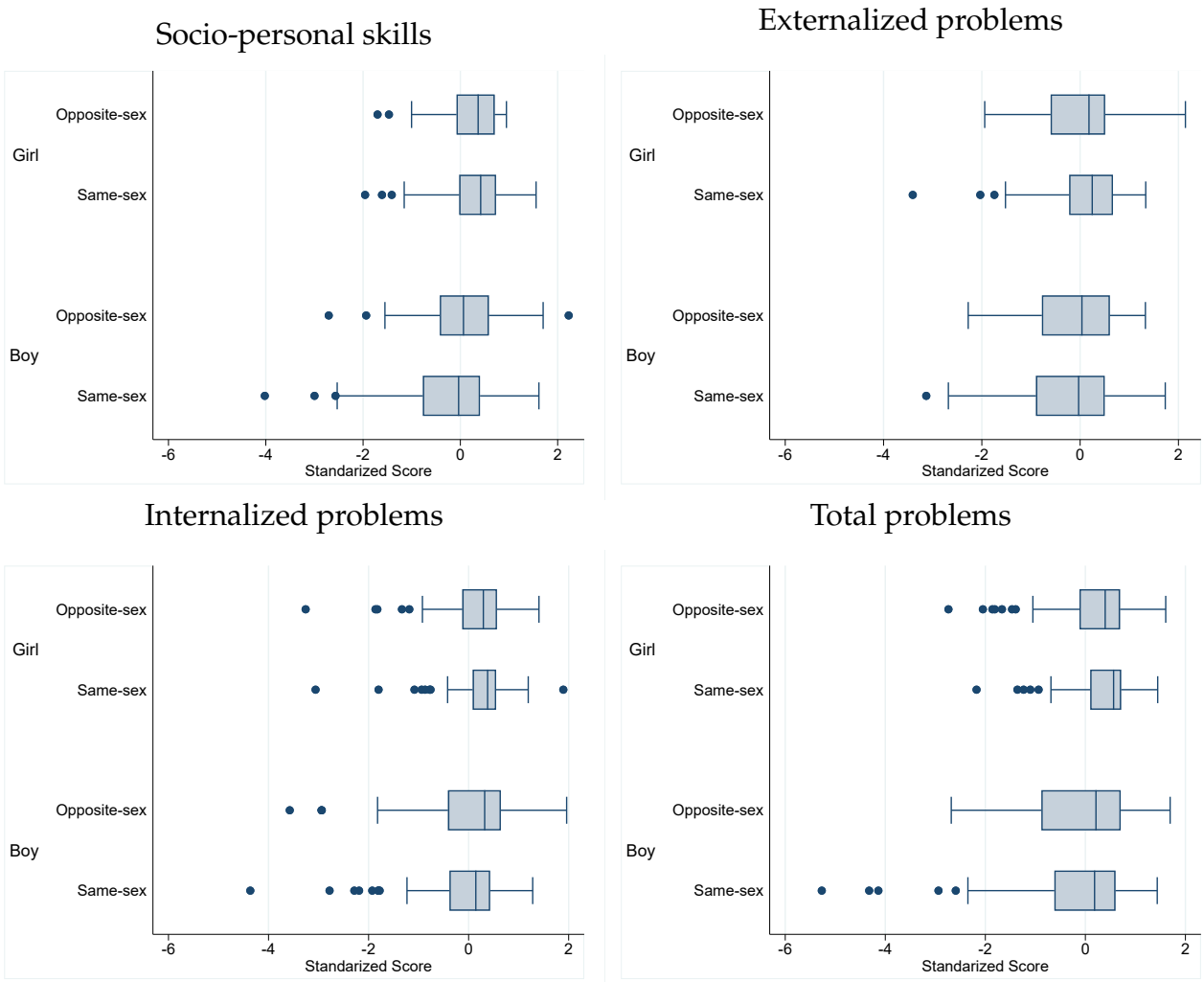


Figure 2.3: Non-Cognitive Skills distributions for first-born children with an opposite-sex or a same-sex second-born sibling, by the sex of the first-born.

2.3.2 Identification Strategy

Recent research on siblings' sex composition shows that the causal effect of a younger sibling's sex can be identified, but not the other way around (Peter et al., 2018; Cools and Patacchini, 2019; Vogl, 2013; Brenøe, 2018).[14] The gender of the older child may affect parents' decision to have another child, which may be correlated with parental preferences for children's sex composition.

[14]In such case, the effect of siblings' sex composition can be biased with parental preferences because the comparison is between girls (boys) with an older brother to girls (boys) with an older sister.

The identification strategy used in this paper is based on the exposure of a first-born child to a same-sex or opposite-sex second-born sibling. This strategy is free from selection bias since the sex of the younger child is random and cannot affect the existence of the older child (Peter et al., 2018). Then, the average treatment effect of having a sibling of the same or opposite sex is identified, as a consequence of the direct exposure of the child to a sister or a brother (Black et al., 2021). The analysis will be conducted separately for girls and boys. Same-sex siblings are defined if the first two siblings are of the same sex, and opposite-sex siblings are defined if they are of opposite sexes. In order to identify the causal effect of sibling's sex on the outcome variables, the main regression equation is the following:

$$y_i = \alpha + \beta d_i + \theta X_i + \eta_i \tag{2.1}$$

where y is the outcome variable of interest, d is an indicator variable taking the value one if a child has a same-sex younger sibling, and X includes individual and family characteristics. Parameter β represents the difference in y for those exposed to same-sex siblings relative to those exposed to opposite-sex siblings.

The data needs to be balanced to ensure there is no selection on observables.[15] Table 2.2 presents the means for children with opposite-sex or same-sex siblings (columns 1 and 2, respectively), the difference between these groups (column 3), t-statistics of this difference (column 4) and the associated p-value (column 5). The upper panel reports result for girls and the lower panel the corresponding for boys. On average, girls and boys who have a younger sibling in the second round are 52 months old, 83% are white, the household size is around 4.5 people, and there is no difference in the per-capita income. The mothers' age in the case of girls is significantly different between those with a same-sex to an opposite-sex younger sibling. Mothers' educational level shows some difference between boys and girls, but not between girls or boys with a same-sex or opposite-sex younger sibling. Around 50% of girls' mothers have not completed high school, and this percentage rises to 60% for the boys' mothers; more than 30% of girls' mothers completed high school and around 15% of boys'; finally, 20% of girls' mothers completed college, reaching 25% in the case of boys' mothers. Birth spacing between siblings is approximately 33 in the case of girls and 34 in the case of boys, and around 33% of girls' siblings and 41% of boys' siblings have less than one year of age.[16] The maximum birth spacing in the data is six years, since children in the

[15]There is no evidence from Uruguay supporting parental preferences over the sex composition of their children, nor selective abortion.

[16]In the full sample of children with a younger sibling, the only exception on the balanced data is the

first round are up to four years old and the second round was performed two years later.

Table 2.2: Summary statistics of first-born children. Variables included in the regression analysis.

	Opposite-sex	Same-sex	Difference	t-stat	p-value
Girls					
Age (in months)	51.2	52.1	0.909	0.716	0.475
Race	0.87	0.78	-0.091	-1.389	0.165
Household size	4.4	4.4	-0.023	-0.107	0.915
Ln per-capita income	8.0	8.3	0.286	1.515	0.132
Region	0.48	0.43	-0.045	-0.516	0.606
Mother's age	27.4	29.5	2.048	1.990	0.049
Mother not completed HS	0.49	0.46	-0.033	-0.384	0.701
Mother completed HS	0.33	0.29	-0.036	-0.448	0.654
Mother completed College	0.18	0.25	0.070	0.979	0.328
Birth spacing (in months)	32.5	33.5	1.082	0.502	0.617
Sibling's less than 1 year old	0.30	0.35	0.052	0.641	0.521
Observations	61	72			
Boys					
Age (in months)	52.9	52.4	-0.489	-0.417	0.677
Race	0.83	0.84	0.017	0.281	0.778
Household size	4.8	4.8	0.071	0.266	0.791
Ln per-capita income	8.0	8.1	0.162	0.999	0.319
Region	0.43	0.40	-0.029	-0.370	0.712
Mother's age	27.9	28.1	0.188	0.181	0.857
Mother not completed HS	0.61	0.60	-0.011	-0.140	0.889
Mother completed HS	0.15	0.14	-0.002	-0.037	0.970
Mother completed College	0.24	0.25	0.013	0.189	0.850
Birth spacing (in months)	36.2	33.2	-3.076	-1.546	0.124
Sibling's less than 1 year old	0.43	0.39	-0.041	-0.524	0.600
Observations	75	83			

Notes: The table reports means and standard deviations of individual and family characteristics. The sample consists of first-born children who have a younger sibling in wave 2.

The variables included as covariates in equation 2.1 are predetermined and do not depend on the sex of the child or the sex of the younger sibling. Table 2.3 shows the probability

birth order of girls, since there are more first-born girls with a same-sex younger sibling (67% to 52%) than second-born girls with a same-sex younger sibling (15% to 27%).

of having a same-sex younger sibling for girls and boys, separately, on the first-born sample and on a sample including all children with a younger sibling, adjusted for the covariates used in the regression analysis, and summarized in Table 2.2.[17][18] As expected, results show that these variables are not significantly correlated with the sex of the younger sibling. Based on the test of joint significance of differences, it cannot be rejected that the characteristics of girls and boys in the same-sex or opposite-sex group are the same (Chi-squared statistic for first-born girls is 9.65 and 10.58 for all girls with a younger sibling; whereas for boys it is 5.15 for first-borns and 4.88 for all boys with a younger sibling). These results support the main identification assumption of the regression analysis.

[17]The omitted category on mother's education is "high school not completed". Birth order is only added when the whole sample of children with at least one younger sibling has been used.

[18]To avoid bad control bias, family size is controlled, the number of siblings living in the household is not, since it is considered an outcome variable itself (Angrist and Pischke, 2008).

Table 2.3: Randomness of the sex of the next younger sibling.

	Next-younger is female (Girls)		Next-younger is male (Boys)	
	First-born	All	First-born	All
Age (in months)	0.0257	0.0034	-0.0077	-0.0044
	(0.024)	(0.017)	(0.022)	(0.016)
Race (1=white)	-0.4816	-0.2077	-0.0379	-0.1442
	(0.486)	(0.332)	(0.464)	(0.326)
Ln Number of people	0.0102	-0.0119	0.0572	0.0333
in the household	(0.152)	(0.114)	(0.125)	(0.089)
Ln per-capita income	0.1960	0.1666	0.2749	0.0709
	(0.238)	(0.202)	(0.216)	(0.170)
Region (1=Montevideo)	-0.3559	0.2537	-0.2928	-0.3667
	(0.380)	(0.289)	(0.355)	(0.267)
Mother's age	0.0822*	0.0205	0.0033	-0.0000
	(0.046)	(0.032)	(0.038)	(0.027)
Mother HS graduate (1=Yes)	-0.6850	-0.3170	0.0764	0.1235
	(0.531)	(0.397)	(0.510)	(0.397)
Mother college graduate (1=Yes)	-0.5091	-0.3492	-0.1797	0.0930
	(0.665)	(0.530)	(0.575)	(0.499)
Birth spacing (in months)	-0.0177	0.0073	-0.0222	-0.0116
	(0.021)	(0.016)	(0.019)	(0.014)
Sibling's less than 1 year old	0.7019	-0.1279	0.1637	-0.0804
(1=Yes)	(0.535)	(0.387)	(0.460)	(0.343)
First child (1=Yes)		0.4091		0.1531
		(0.527)		(0.477)
Second child (1=Yes)		-0.4364		0.3911
		(0.531)		(0.468)
Constant	-3.9065	-2.2749	-1.1897	0.1175
	(2.425)	(1.856)	(2.034)	(1.745)
Test of joint significance - Chi2	9.65	10.58	5.15	4.88
p-value	0.47	0.57	0.88	0.96
Observations	133	224	158	260

Notes: The table reports logit regression on the sex of the next younger sibling. The last row reports Chi-squared and *p-value* of tests of significance of all of the variables included in the regression.
Robust standard errors (in parenthesis).
Standard errors adjusted for clustering at the family level (in parenthesis).
Estimates marked *** are significant at 1%, ** at 5%, * at 10%.

2.4 Results

This section presents, firstly, the effects of having a younger same-sex or opposite-sex sibling on motor skills, cognitive and non-cognitive development, and secondly, the potential channels that could explain the differences found.

2.4.1 Childhood Development

The middle-term effects of siblings' sex composition are estimated using the equation (2.1) for each outcome variable of standardized scores of ASQ-3 and reverse scales of CBCL tests. The inference is based on standard errors and *p-values* adjusted for multiple hypotheses testing, using the step-down procedure described in Romano and Wolf (2005, 2016).[19] This paper takes into account three groups of hypotheses: the "Motor Skills" hypothesis includes the measures of gross motor skills and fine motor skills, measured by the ASQ-3 test; the "Cognitive" hypothesis includes communication and problem-solving skills from the ASQ-3 test; and the "Non-Cogninitive Skills" hypothesis group analyses the socio-personal component of the ASQ-3 test, and the three measures of CBCL test. Each one of these hypotheses represents a group being jointly tested.

Table 2.4 presents the results for children with a same-sex second-born sibling compared to those with an opposite-sex second-born sibling (parameter β), for girls and boys (panels A and B, respectively).[20] Girls have no differential effect on their motor skills, cognitive or non-cognitive development resulting from having a same-sex or opposite-sex younger sibling. In turn, boys have significant negative effects from having a same-sex younger sibling, compared to boys with an opposite-sex younger sibling, in motor skills and non-cognitive development.

For boys, the average difference between having a younger brother and a younger sister, in terms of gross-motor and fine-motor skills, is of -0.28 standard deviations and -0.51 standard deviations, respectively. Regarding non-cognitive skills, boys with a younger brother have less socio-personal abilities, and higher levels of internalized and total problems.[21]

[19] Romano and Wolf (2005) have developed a procedure to avoid false rejection of true null hypotheses when multiple outcomes are jointly tested. This procedure takes into account the dependence structure of the test statistics by using resampling methods from the original data, making it more powerful than the Bonferroni or Holm corrections.

[20] All the tables in this section present estimates on panels A.1 and B.1 without covariates and on panels A.2 and B.2 with the control variables described in the previous section. Results when covariates are included improve the precision of the estimate but do not change the main results of the treatment variable.

[21] The opposite of the standardized test is used in the presentation of test results associated with the CBCL.

The socio-personal skills are 0.29 standard deviations lower for boys who grow-up with a same-sex second-born sibling than those with a second-born sister; whereas internalized and total problems score 0.40 and 0.30 standard deviations lower, respectively. When these hypotheses were adjusted for multiple hypotheses testing, only the socio-personal component loses its significance.[22]

Results indicate that sibling's gender has a negative impact on boys' motor skills and non-cognitive development, but does not affect girls. These results are in line with the results obtained by Cyron et al. (2017), who found that boys had strong negative issues caused by having a younger brother, but no significant effects on girls. The next section presents potential mechanisms that could explain these results.

However, these aspects will still be discussed as problems associated with the development of non-cognitive skills.

[22]Table 2.A.3 presents the results of the sample that includes all children with a younger sibling. Boys who have a same-sex younger sibling have lower levels of fine motor skills and higher levels of internalized and total problems. Results adjusted for multiple hypotheses testing show that only internalized problems are significant at 10%.

Table 2.4: Childhood Development Outcomes

	ASQ-3					*CBCL*		
	Motor Skills[a]		*Cognitive Skills*[b]			*Non-Cognitive Skills*[c]		
	Gross-motor	Fine-motor	Commu-nication	Problem solving	Socio-personal	Externa-lized	Interna-lized	Total problems
Panel A: Girls								
Opposite-sex mean	0.0716	0.2568	0.1206	0.1350	0.2147	-0.0356	0.0851	0.0309
	(0.069)	(0.107)	(0.103)	(0.112)	(0.081)	(0.112)	(0.088)	(0.098)
Panel A.1: No controls								
Same-sex	-0.0137	0.2935*	0.1099	0.2296*	0.0489	0.1737	-0.0817	0.0184
	(0.114)	(0.152)	(0.131)	(0.138)	(0.113)	(0.152)	(0.135)	(0.138)
	[0.908]	[0.124]	[0.393]	[0.172]	[0.9002]	[0.535]	[0.832]	[0.9002]
Prob>F	0.904	0.058	0.404	0.099	0.668	0.256	0.546	0.894
R2	0.000	0.046	0.005	0.022	0.001	0.010	0.003	0.000
Panel A.2: Individual and family characteristics								
Same-sex	-0.0575	0.2367	0.0227	0.1334	-0.0076	0.0620	-0.2054	-0.0919
	(0.117)	(0.180)	(0.136)	(0.139)	(0.117)	(0.147)	(0.135)	(0.136)
	[0.643]	[0.208]	[0.866]	[0.515]	[0.932]	[0.892]	[0.277]	[0.798]
Prob>F	0.059	0.003	0.066	0.000	0.379	0.000	0.000	0.000
R2	0.106	0.249	0.149	0.189	0.071	0.258	0.244	0.250
Observations	133	79	133	133	133	133	133	133
Panel B: Boys								
Opposite-sex mean	0.0996	-0.0287	0.0383	-0.0179	0.0323	-0.1277	0.0944	0.0203
	(0.088)	(0.123)	(0.117)	(0.119)	(0.098)	(0.108)	(0.081)	(0.089)
Panel B.1: No controls								
Same-sex	-0.2474*	-0.4350**	-0.1052	-0.1376	-0.3110**	-0.0810	-0.3568***	-0.2417*
	(0.146)	(0.209)	(0.157)	(0.179)	(0.147)	(0.155)	(0.128)	(0.135)
	[0.11]	[0.11]	[0.661]	[0.661]	[0.084]	[0.603]	[0.01]	[0.106]
Prob>F	0.094	0.040	0.505	0.444	0.036	0.604	0.006	0.076
R2	0.017	0.040	0.003	0.004	0.027	0.002	0.046	0.020
Panel B.2: Individual and family characteristics								
Same-sex	-0.2797*	-0.5086***	-0.1434	-0.1533	-0.2908**	-0.1458	-0.4014***	-0.3068**
	(0.143)	(0.195)	(0.157)	(0.174)	(0.147)	(0.149)	(0.127)	(0.129)
	[0.066]	[0.028]	[0.567]	[0.567]	[0.118]	[0.321]	[0.008]	[0.042]
Prob>F	0.043	0.000	0.149	0.025	0.315	0.002	0.000	0.000
R2	0.101	0.298	0.076	0.126	0.068	0.132	0.184	0.174
Observations	158	95	158	158	158	158	158	158

Notes: The table reports the average treatment effect of having a same-sex second-born sibling, for separate regressions of ASQ-3 and CBCL tests, by the sex of the first-born child. Panels A.1 and B.1 do not have controls. Panels A.2 and B.2 include birth spacing (in months) with the second-born child; a dummy variable equals to one when the sibling is less than one year old; age in months of the child; race (white=1); region (Montevideo=1), the logarithm of people living in the household; the logarithm of per capita income; mother's age; a dummy variable if the mother completed high school, and a dummy variable if the mother is college graduate. Robust standard errors (in parenthesis). Estimates marked *** are significant at 1%, ** at 5%, * at 10%.

[a] Romano-Wolf *p-values* (in square brackets) for motor skills (two coefficients).

[b] Romano-Wolf *p-values* (in square brackets) for cognitive skills (two coefficients).

[c] Romano-Wolf *p-values* (in square brackets) for non-cognitive skills (four coefficients).

2.4.2 Potential Mechanisms

This subsection analyzes all the potential mechanisms, discussed in section 2.2.2, that could explain the differences observed as a result of having a same-sex younger sibling, his or her influence on parental behavior regarding household structure, investments on active time, expectations, attendance to school, health care, parenting styles, and household environment.[23] The focus of this paper is on the middle-term outcomes, thus the direct sibling-to-sibling interactions are assumed to be negligible, as well as other attitudes or behaviors of the parents, like intergenerational transmission of gender norms. Economies of scale are not studied, given that the effect of having a same-sex sibling, if any, should be positive.

2.4.2.1 Family Structure

Parents' attitudes toward their children can be different depending on their sex composition. In this subsection, different aspects of family structure, such as living with both biological parents, the number of siblings living in the household, and mothers working status, are analyzed to determine if there is a factor that can account for the differences observed. Table 2.5 reports the results of the linear estimation using equation (2.1) on each one of these possible mechanisms.

On average, 72% of first-born girls who have an opposite-sex second-born sibling, live with both parents. This percentage is not statistically different in the case of girls with a same-sex younger sibling. On the other hand, 76% of first-born boys who have an opposite-sex younger sibling live with both parents, but this probability falls by almost 19 percentage points if they have a same-sex younger sibling. It is worth noting that only one of the first-born boys of the sample lives only with his father, and none of the first-born girls live only with their father. Therefore, the circumstance of not living with both parents can be associated with the absence of the biological father on child-rearing.[24] Results show that the probability of living with both parents in mixed-sex sibling composition is roughly the same for girls and boys (0.72 for first-born girls and 0.76 for first-born boys). However,

[23]The inference is based on robust standard errors in the estimations. Multiple hypotheses testing is not valid since each hypothesis test corresponds to a single aspect of parental behavior.

[24]Results in the short term show that the probability of living with both parents is not affected by the sex of the younger child, neither for girls nor boys. To explore some explanation, I estimate the effect of having a newborn sibling, estimated on all children and then specifically on those who are first-born children in the first round, and then matched to the probability of living with both parents. Results indicate that parents who do not experience a second birth have a higher likelihood of living apart. Therefor, the short-term effect of having a new child is keeping the family together.

the difference is significant when considering the sex of the younger sibling (0.75 for first-born girls with a same-sex sibling versus 0.57 for first-born boys with a same-sex younger sibling).[25]

The number of siblings living in the household is around one, for girls and boys, and there are no significant differences between children who have a same-sex or opposite-sex younger sibling. This hypothesis needs to be tested on mothers who are not in their reproductive age, which is the case of the sample in this study.[26] Around 55% of mothers in the first-born sample worked, and there is no significant difference between those who have same-sex or mixed-sex children.

[25]There is no difference in the probability of having the same father among girls or boys. This result is not shown in tables.

[26]Table 2.A.4 of the Appendix presents the overall estimates for children with a younger sibling in the sample. Results have the same sign and significance than those presented for first-born.

Table 2.5: Family structure

	Two-parent household	Number of siblings in household	Mother worked
Panel A: Girls			
Opposite-sex mean	0.7213	1.0656	0.5574
	(0.057)	(0.031)	(0.064)
Panel A.1: No controls			
Same-sex	0.0842	0.0316	0.0954
	(0.074)	(0.047)	(0.085)
Prob>F	0.260	0.506	0.266
R2	0.009	0.003	0.009
Panel A.2: Individual and family characteristics			
Same-sex	0.0346	0.0247	-0.0027
	(0.061)	(0.045)	(0.069)
Prob>F	0.000	0.145	0.000
R2	0.338	0.200	0.445
Observations	133	133	133
Panel B: Boys			
Opposite-sex mean	0.7600	1.0667	0.5467
	(0.049)	(0.028)	(0.057)
Panel B.1: No controls			
Same-sex	-0.1334*	0.0297	0.0557
	(0.072)	(0.043)	(0.078)
Prob>F	0.069	0.497	0.482
R2	0.021	0.003	0.003
Panel B.2: Individual and family characteristics			
Same-sex	-0.1890**	0.02376	0.0365
	(0.080)	(0.041)	(0.059)
Prob>F	0.000	0.116	0.000
R2	0.314	0.191	0.400
Observations	158	158	158

Notes: The table reports the average treatment effect of having a same-sex second-born sibling for separate regressions of family characteristics, by the sex of the first-born child. The effects relating to the binary variables have been estimated using a logit regression model. The reported coefficients are marginal effects evaluated at the average probability and then averaged over the sample. Panels A.1 and B.1 do not have controls. Panels A.2 and B.2 include birth spacing (in months) with the second-born child; a dummy variable equals to one if the sibling is less than one year old; age in months of the child; race (white=1); region (Montevideo=1), logarithm of people living in the household; the logarithm of per capita income; mother's age; a dummy variable if the mother completed high school, and a dummy variable if the mother is college graduate. Robust standard errors (in parenthesis). Estimates marked *** are significant at 1%, ** at 5%, * at 10%.

2.4.2.2 Parental Investment and Expectations

This subsection analyzes parents' behavior towards their first-born child using variables related to quality time spent with their children, decisions regarding their education, and decisions about their vaccination; as well as expectations related to their children's educational achievement. Results are presented in Table 2.6.

To approximate an index of parental quality time investment, a principal component analysis was conducted. The variables used to construct this index are proxies of active interactions between parents and their children (e.g. whether parents told stories, taught new games, or sang to them last week).[27] The time parents spend with their first-born son or daughter is slightly above the median, and there are no significant differences with the time parents spend with their first-born daughters if they have a second child, girl or boy. Conversely, parents of first-born boys spend less active time with them if they have a second son than if they have a daughter (-0.3 standard deviation lower).

This result can be caused by a reinforcement parental behavior, or a complementarity effect, if they sing, play, or tell stories more often to girls than boys. If this is the case, results on the difference between active time parents spend with girls (boys) who have a same-sex younger sibling, and those who have an opposite-sex younger sibling should be positive (negative) and significant. Although this difference is present in the boys' sample (a negative and significant difference), it is not accurate in the girls' sample. A case parents of mixed-sex children composition spend less time with the oldest child to spend more time with the youngest could be a substitution effect. For example, if parents need to learn how to raise their younger child. In this case, the effect of having a same-sex younger sibling should be positive, irrespective of the sex of the oldest child.

An other explanation for this result is related to the reverse gender-gap hypothesis. If parents spend more quality time with girls, first-born boys could benefit from a positive externality of having a younger sister, relative to those having a younger brother.

To test if parents invest differently in the formal education of their children, attendance to a formal school was analyzed. On average, 89% of girls with an opposite-sex younger sibling attend school, while this percentage increases to 99% in the case of boys. Parents do not behave differently towards their first-born daughters when they have a second child, boy or girl. Again, there are significant differences among first-born boys, where the probability of attending school falls by almost 18 percentage points when first-born boys have a same-sex sibling. This result could also be associated with the reverse gender-gap hypoth-

[27] The entire sample is used to construct this index.

esis, which leads to economies of scale. Parents supporting this hypothesis may decide to enroll their daughters sooner to a formal school, and to reach some economies of scale, they decide to send both into school.

Parents may also have different expectations about school achievements from their sons and daughters; 67% of girls' parents and 75% boy's parents with a second opposite-sex child expect their children to attend college. The differences regarding same-sex to opposite-sex siblings are not significant among girls or boys.

As a measure of health investment, I test if parents' behavior is tested in terms of flu immunization of their sons and daughters. This kind of vaccine is not compulsory, although doctors recommend it. A different prevalence of this attitude towards one of the sexes would show some relative preference from parents. Results show an average rate of vaccination of 26%, with no difference between boys or girls, regardless of the sex of the younger sibling.

Table 2.6: Parental investment and expectations

	Parental investment			*Health-care*
	Active Time	Attending School	College Expectations	Flu immunization
Panel A: Girls				
Opposite-sex mean	0.0613	0.8852	0.6721	0.2787
	(0.121)	(0.041)	(0.060)	(0.057)
Panel A.1: No controls				
Same-sex	0.1697	0.0175	0.03620	0.0824
	(0.145)	(0.053)	(0.080)	(0.080)
Prob>F	0.246	0.746	0.656	0.312
R2	0.010	0.000	0.001	0.007
Panel A.2: Individual and family characteristics				
Same-sex	0.1595	-0.0034	-0.0282	0.0019
	(0.144)	(0.043)	(0.068)	(0.080)
Prob>F	0.000	0.088	0.000	0.000
R2	0.208	0.193	0.297	0.221
Observations	133	133	133	133
Panel B: Boys				
Opposite-sex mean	0.0798	0.9867	0.7467	0.2533
	(0.102)	(0.013)	(0.050)	(0.050)
Panel B.1: No controls				
Same-sex	-0.2954*	-0.1071***	-0.0840	0.0237
	(0.167)	(0.038)	(0.072)	(0.070)
Prob>F	0.080	0.006	0.249	0.737
R2	0.019	0.044	0.008	0.001
Panel B.2: Individual and family characteristics				
Same-sex	-0.3021*	-0.1769***	-0.0899	0.0262
	(0.170)	(0.059)	(0.060)	(0.060)
Prob>F	0.010	0.173	0.000	0.000
R2	0.139	0.201	0.306	0.251
Observations	158	158	158	158

Notes: Robust standard errors (in parenthesis). Estimates marked *** are significant at 1%, ** at 5%, * at 10%. Other notes see Table 2.5.

2.4.2.3 Parental Practices

This section approximates parental practices through two measures of parenting styles, authoritarian and authoritative, and two measures of home environment index, warm and harsh environments.[28] This paper follows the proposal made by Perazzo et al. (2018), who constructed two composite indexes of parental styles through principal component analysis. The authoritarian style is based on obedience, imposing parents' choices on their children and sometimes implying the use of punishment to regulate children's behavior. The permissive parenting style gives children independence to choose and does not impose almost any rule. In the authoritative, parents set some rules, but these do not aim to impose their will, but to promote their development. It is associated with effective involvement, active control of children's activities, and responsiveness to their demands, combined with strategies of non-violent discipline (Perazzo et al., 2018). These dimensions have been standardized.

The second round of the ENDIS collects items corresponding to the HOME observed environment. These items can be grouped in the subscales of receptivity (warmth index, when the caregiver speaks to the child with affection, talks with the child spontaneously, expresses sensitivity, and is responsive with to the child) or punishment (harshness index, the caregiver yells at the child or beats him or her during the interview). A higher score in the warmth -harshness- subscale indicates that the child is exposed to better -worse- parental practices.[29]

The first and second columns of Table 2.7 present the results of authoritarian and authoritative parenting styles, and the third and fourth columns the results of the warmth and harshness indexes. Parents do not have different behaviors regarding girls or boys who have a same-sex or opposite-sex sibling. These results may indicate that parents have no reinforcement in their parental practices. If parents have a reinforcement behavior, we may observe a positive (negative) effect in the authoritarian style and harshness indexes for boys (girls), and a negative (positive) effect in the authoritative and warmth indexes.

[28] The measures of home environment index are taken from the Home Observation Measurement of the Environment (HOME) questionnaire applied on the survey.

[29] Details on the construction of these indices can be found in Bando et al. (2016).

Table 2.7: Parental practices

	Parenting styles		*Home environment*	
	Authoritarian	Authoritative	Warm	Harsh
Panel A: Girls				
Opposite-sex mean	-0.3100	0.0489	1.3934	0.4754
	(0.115)	(0.098)	(0.172)	(0.127)
Panel A.1: No controls				
Same-sex	0.1921	-0.1184	-0.060	-0.0170
	(0.156)	(0.163)	(0.231)	(0.157)
Prob>F	0.223	0.468	0.795	0.913
R2	0.011	0.003	0.000	9.340
Panel A.2: Individual and family characteristics				
Same-sex	0.2496	-0.1948	0.0836	0.0929
	(0.168)	(0.159)	(0.249)	(0.151)
Prob>F	0.050	0.062	0.031	0.048
R2	0.133	0.102	0.104	0.101
Observations	133	133	133	133
Panel B: Boys				
Opposite-sex mean	-0.0498	0.1259	1.8133	0.8667
	(0.098)	(0.080)	(0.182)	(0.139)
Panel B.1: No controls				
Same-sex	-0.061	-0.005	-0.0904	-0.1196
	(0.144)	(0.119)	(0.244)	(0.195)
Prob>F	0.674	0.964	0.712	0.542
R2	0.001	0.000	0.001	0.002
Panel B.2: Individual and family characteristics				
Same-sex	-0.032	-0.003	-0.015	-0.1112
	(0.146)	(0.127)	(0.232)	(0.189)
Prob>F	0.161	0.884	0.000	0.056
R2	0.071	0.032	0.192	0.163
Observations	158	158	158	158

Notes: Robust standard errors (in parenthesis). Estimates marked *** are significant at 1%, ** at 5%, * at 10%. Other notes see Table 2.5.

2.4.3 Importance of Mechanisms

The negative effects observed in first-born boys who have a second-born brother, compared to those who have a second-born sister, are mainly explained by the decrease in quality time parents spend with them, as well as the lower probability of living with both parents and attending school. To analyze the importance of these mechanisms on the different components of motor skills, cognitive and non-cognitive development, these channels were added, one at a time, to the main regression analysis. Results for first-born boys are shown in Table 2.8.[30] Panel A shows the results regarding family structure, Panel B shows the results for parental investment in quality time, Panel C shows the results for parental investment in school attendance, and Panel D adds up all these variables to examine whether the main effects remain.

Living with both parents improves the performance of fine-motor skills, socio-personal skills, and reduces externalized and total problems. Parental time investment is negatively correlated with gross-motor skills, but it is positively correlated with having less externalized and total problems.[31] Finally, school attendance enhances the performance of gross-motor skills.

When all these channels are added in the main estimation, negative effects on boys as a result of having a same-sex younger sibling remain significant on fine motor skills and internalized problems (after correcting by the Romano-Wolf multiple hypotheses testing), and the mechanism that turns out to be most critical in reducing these differences is living with both parents.

[30] Results for girls are not significant, and therefore are not shown.

[31] The estimation over the CBCL tests are presented with the opposite value of the standardized test.

Table 2.8: Estimation of main mechanisms affecting motor skills, cognitive, and non-cognitive development (first-born boys subsample)

	ASQ-3					*CBCL*		
	Motor Skills[a]		*Cognitive Skills*[b]			*Non-Cognitive Skills*[c]		
	Gross-motor	Fine-motor	Commu-nication	Problem solving	Socio-personal	Externa-lized	Interna-lized	Total problems
Panel A: biological parents in the household								
Same-sex	-0.2460*	-0.4356**	-0.1247	-0.1092	-0.2390*	-0.0871	-0.3722***	-0.2611**
	(0.144)	(0.183)	(0.150)	(0.168)	(0.144)	(0.146)	(0.122)	(0.123)
	[0.107]	[0.073]	[0.674]	[0.674]	[0.223]	[0.564]	[0.015]	[0.117]
Two-parent household	0.2430	0.5979**	0.1344	0.3174	0.3725**	0.4224**	0.2101	0.3290*
	(0.186)	(0.229)	(0.194)	(0.216)	(0.171)	(0.203)	(0.182)	(0.178)
Prob>F	0.1076	0.0002	0.2345	0.0434	0.1652	0.0003	0.0003	0.0000
R2	0.111	0.337	0.078	0.137	0.091	0.160	0.193	0.195
Penal B: parental investment								
Same-sex	-0.3162**	-0.5066**	-0.1364	-0.1282	-0.2606*	-0.0908	-0.3905***	-0.2712**
	(0.146)	(0.202)	(0.157)	(0.176)	(0.148)	(0.147)	(0.127)	(0.127)
	[0.033]	[0.033]	[0.606]	[0.606]	[0.173]	[0.534]	[0.003]	[0.103]
Active time	-0.1204**	0.0078	0.0233	0.0834	0.0999	0.1823**	0.0362	0.1178*
	(0.058)	(0.098)	(0.075)	(0.085)	(0.075)	(0.083)	(0.066)	(0.068)
Prob>F	0.0270	0.0003	0.2384	0.0294	0.3338	0.0008	0.0003	0.0000
R2	0.117	0.298	0.076	0.131	0.079	0.167	0.185	0.193
Panel C: school attendance								
Same-sex	-0.1638	-0.5215**	-0.0722	-0.0843	-0.2290	-0.1088	-0.3731***	-0.2725**
	(0.139)	(0.201)	(0.154)	(0.155)	(0.147)	(0.151)	(0.126)	(0.125)
	[0.279]	[0.039]	[0.848]	[0.848]	[0.259]	[0.506]	[0.011]	[0.091]
Attending school	1.1365**	-0.0709	0.6980	0.6763	0.6058	0.3629	0.2777	0.3367
	(0.511)	(0.466)	(0.504)	(0.705)	(0.479)	(0.456)	(0.389)	(0.433)
Prob>F	0.0148	0.0000	0.2055	0.0356	0.3030	0.0039	0.0006	0.0001
R2	0.177	0.298	0.102	0.144	0.089	0.139	0.189	0.182
Panel D: all relevant variables								
Same-sex	-0.1811	-0.4799**	-0.0559	-0.0340	-0.1686	-0.0190	-0.3447***	-0.2098*
	(0.146)	(0.196)	(0.149)	(0.156)	(0.147)	(0.147)	(0.124)	(0.121)
	[0.257]	[0.053]	[0.916]	[0.916]	[0.475]	[0.914]	[0.033]	[0.277]
Two-parent household	0.1976	0.6413**	0.0493	0.2140	0.2721*	0.3067	0.1726	0.2465
	(0.196)	(0.249)	(0.188)	(0.221)	(0.163)	(0.200)	(0.176)	(0.172)
Active time	-0.1162*	-0.0332	0.0322	0.0792	0.0899	0.1657*	0.0278	0.1051
	(0.060)	(0.102)	(0.072)	(0.085)	(0.075)	(0.086)	(0.066)	(0.069)
Attending school	1.0409**	-0.2255	0.6953	0.6434	0.5603	0.3338	0.2387	0.3037
	(0.522)	(0.465)	(0.507)	(0.705)	(0.470)	(0.425)	(0.376)	(0.405)
Prob>F	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
R2	0.194	0.341	0.103	0.156	0.115	0.190	0.198	0.214
Observations	158	95	158	158	158	158	158	158

Notes: The table reports the coefficient of the same-sex dummy of first-born children and the main mechanisms, added one at a time, for separate regressions of ASQ-3 and CBCL tests. Panels A, B and C include birth spacing (in months) with the second-born sibling; a dummy variable equals to one if the sibling is less than one year old; age in months of the child; race (white=1); region (Montevideo=1); the logarithm of people living in the household; the logarithm of per-capita income, mother's age; a dummy variable if the mother completed high school, and a dummy variable if the mother is college graduate. Robust standard errors (in parenthesis). Estimates marked *** are significant at 1%, ** at 5%, * at 10%.

[a] Romano-Wolf *p-values* (in square brackets) for motor skills (two coefficients).

[b] Romano-Wolf *p-values* (in square brackets) for cognitive skills (two coefficients).

[c] Romano-Wolf *p-values* (in square brackets) for non-cognitive skills (four coefficients).

2.5 Conclusions

This paper evaluates the impact of siblings' sex composition on motor skills, cognitive and non-cognitive development of children. The main contribution of this paper is to disentangle the mechanisms behind the results found and to add to the growing literature on siblings' sex composition and child-gender. The data used is the Uruguayan ENDIS panel, which collects information on family and socioeconomic characteristics, as well as detailed information on nutrition and cognitive and non-cognitive development. In order to identify the siblings' sex composition effect, the sample was restricted to first-born boys and girls with at least one younger sibling. Children in the sample are between two and six years old.

Results suggest that in the case of boys, having a same-sex younger sibling is associated with lower motor-skills and non-cognitive development, while in the case of girls, a same-sex younger sibling does not affect her performance in the analyzed dimensions. The analyzed channels are related to changes in parental behavior depending on the sex composition of their children, such as changes in family structure, parental investment, expectations and parental practices. The economies of scale hypothesis was not analyzed, because the main effect should be positive, not negative. The "peer-to-peer" effect between siblings was not analyzed either because, in this stage of life, the most influential channel has to do with parental behavior.

The importance of the evidence presented in this paper is because it shows that parental behavior and investments are many times affected by the sex composition of children. More importantly, these differences seem to affect childhood development in some of the most important stages of life.

Among boys, the negative effects found in motor-skills and non-cognitive development are mainly explained by the absence of the father in the household. These findings are in line with previous studies, which found significant consequences of father absence in the non-cognitive development dimension (McLanahan et al., 2013; Bertrand and Pan, 2013).

The fact that parents spend less active time with their first-born sons when they have a younger brother, compared to those who have a younger sister, is in line with a complementarity effect and the reverse gender gap hypothesis: boys with a younger sister could benefit from parents spending more quality time with younger girls (Bertrand and Pan, 2013; Baker and Milligan, 2016; Autor et al., 2019). Assuming parents raise their children together, if they were to choose to spend more quality time with one of their children, they would probably end up spending more time with both. Thus, having a younger sister is a

positive externality for boys.

Finally, first-born boys who have a younger brother are less likely to attend school. This can also be related to the reverse gender gap hypothesis and the existence of some economies of scale. If the second-born is female, then parents who support the reverse gender gap may decide for their daughters to attend school sooner than they would decide for their sons. If none of the children attend school, then childcare takes place in the household. If one of the children is in school, the better solution to avoid hiring a nanny or staying at home with a relative is to send the other child to school too.

According to the evidence presented, living with both biological parents is positively correlated with improved fine-motor and socio-personal skills, and negatively correlated with increased externalized and total problems. Investments in quality time have a positive correlation with communication, problem-solving, and socio-personal skills, and a negative correlation with an increase in externalized problems. Finally, attending school has a positive correlation with improved gross motor skills.

Bibliography

Angrist, J. and Evans, W. (1998). Children and their Parents' Labor Supply: Evidence from exogenous variation in family size. *American Economic Review*, 88(3):450–77.

Angrist, J., Lavy, V., and Schlosser, A. (2010). Multiple Experiments For The Causal Link Between The Quantity And Quality Of Children. *Journal of Labor Economics*, 28(4):773–824.

Angrist, J. D. and Pischke, J.-S. (2008). *Mostly Harmless Econometrics: An empiricist's companion*. Princeton university press.

Åslund, O. and Grönqvist, H. (2010). Family Size and Child Outcomes: Is there really no trade-off? *Labour Economics*, 17(1):130–139.

Autor, D., Figlio, D., Karbownik, K., Roth, J., and Wasserman, M. (2019). Family Disadvantage and the Gender Gap in Behavioral and Educational Outcomes. *American Economic Journal: Applied Economics*, 11(3):338–81.

Baker, M. and Milligan, K. (2016). Boy-girl Differences in Parental Time Investments: Evidence from three countries. *Journal of Human capital*, 10(4):399–441.

Bando, R., López-Boo, F., and Li, X. (2016). Sex-differences in Language and Socio-emotional Skills in Early Childhood. IDB Working Papers Series Nro., IDB–WP–714.

Becker, G. S. and Lewis, H. G. (1973). On the Interaction Between the Quantity and Quality of Children. *Journal of Political Economy*, 81(2):S279–S288.

Becker, G. S. and Tomes, N. (1976). Child Endowments and the Quantity and Quality of Children. *Journal of political Economy*, 84(4, Part 2):S143–S162.

Bertrand, M. and Pan, J. (2013). The Trouble with Boys: Social Influences and the Gender Gap in Disruptive Behavior. *American Economic Journal: Applied Economics*, 5(1):32–64.

Black, S. E., Breining, S., Figlio, D. N., Guryan, J., Karbownik, K., Nielsen, H. S., Roth, J., and Simonsen, M. (2021). Sibling Spillovers. *The Economic Journal*, 131(633):101–128.

Black, S. E., Devereux, P. J., and Salvanes, K. G. (2005). The more the merrier? the effect of family size and birth order on children's education. *The Quarterly Journal of Economics*, 120(2):669–700.

Black, S. E., Devereux, P. J., and Salvanes, K. G. (2010). Small family, smart family? family size and the iq scores of young men. *Journal of Human Resources*, 45(1):33–58.

Bono, E. D., Francesconi, M., Kelly, Y., and Sacker, A. (2016). Early maternal time investment and early child outcomes. *The Economic Journal*, 126(596):F96–F135.

Brenøe, A. A. (2018). Origins of Gender Norms: Sibling Gender Composition and Women's Choice of Occupation and Partner. Technical report, Department of Economics-University of Zurich.

Brenøe, A. A. and Lundberg, S. (2018). Gender Gaps in the Effects of Childhood Family Environment: Do they persist into Adulthood? *European Economic Review*, 109:42–62.

Butcher, K. F. and Case, A. (1994). The Effect of Sibling Sex Composition on Women's Education and Earnings. *The Quarterly Journal of Economics*, 109(3):531–563.

Carneiro, P., Crawford, C., and Goodman, A. (2007). The Impact of Early Cognitive and Non-cognitive Skills on Later Outcomes. Centre for the Economics of Education, London School of Economics.

Conti, G. and Heckman, J. J. (2012). The Economics of Child Well-Being. Working Paper 18466, National Bureau of Economic Research.

Cools, A. and Patacchini, E. (2019). The Brother Earnings Penalty. *Labour Economics*, 58:37–51.

Cunha, F. and Heckman, J. (2007). The technology of skill formation. *The American Economic Review*, 97(2):31–47.

Currie, J. and Almond, D. (2011). Human Capital Development Before Age Five. In *Handbook of labor economics*, Volume 4, Pages 1315–1486. Elsevier.

Cyron, L., Schwerdt, G., and Viarengo, M. (2017). The Effect of Opposite Sex Siblings on Cognitive and Noncognitive Skills in Early Childhood. *Applied Economics Letters*, 24(19):1369–1373.

Dahl, G. B. and Moretti, E. (2008). The Demand for Sons. *The Review of Economic Studies*, 75(4):1085–1120.

Del Boca, D., Flinn, C., and Wiswall, M. (2013). Household Choices and Child Development. *Review of Economic Studies*, 81(1):137–185.

Depew, B. and Price, J. (2018). Marriage and the Economic Status of Women with Children. *Review of Economics of the Household*, 16(4):1049–1061.

Doepke, M., Sorrenti, G., and Zilibotti, F. (2019). The Economics of Parenting. *Annual Review of Economics*, 11:55–84.

Doepke, M. and Zilibotti, F. (2017). Parenting with Style: Altruism and Paternalism in Intergenerational Preference Transmission. *Econometrica*, 85(5):1331–1371.

Fan, X., Fang, H., and Markussen, S. (2015). Mothers' Employment and Children's Educational Gender Gap. Technical report, National Bureau of Economic Research, WP–21183.

Fan, Y., Yi, J., Yuan, Y., and Zhang, J. (2018). The Glorified Mothers of Sons: Evidence from Child Sex Composition and Parental Time Allocation in Rural China. *Journal of Economic Behavior & Organization*, 145:249–260.

Fiorini, M. and Keane, M. P. (2014). How the Allocation of Children's Time Affects Cognitive and Noncognitive Development. *Journal of Labor Economics*, 32(4):787–836.

Fortin, N. M., Oreopoulos, P., and Phipps, S. (2015). Leaving Boys Behind Gender Disparities in High Academic Achievement. *Journal of Human Resources*, 50(3):549–579.

Ginja, R., Jans, J., and Karimi, A. (2017). Parental Investments in Early Life and Child Outcomes Evidence from Swedish Parental Leave Rules. Available at SSRN: `https://ssrn.com/abstract=3069449` or `http://dx.doi.org/10.2139/ssrn.3069449`

Golding, P. and Fitzgerald, H. E. (2019). The Early Biopsychosocial Development of Boys and the Origins of Violence in Males. *Infant mental health journal*, 40(1):5–22.

Guo, R., Yi, J., and Zhang, J. (2017). Family Size, Birth Order, and Tests of the Quantity–Quality model. *Journal of Comparative Economics*, 45(2):219–224.

Heckman, J., Pinto, R., and Savelyev, P. (2013). Understanding the Mechanisms Through which an Influential Early Childhood Program Boosted Adult Outcomes. *American Economic Review*, 103(6):2052–86.

Heckman, J. J. (2008). Schools, Skills, and Synapses. *Economic inquiry*, 46(3):289–324.

Heckman, J. J. and Rubinstein, Y. (2001). The Importance of Noncognitive Skills: Lessons from the GED Testing Program. *American Economic Review*, 91(2):145–149.

Juhn, C., Rubinstein, Y., and Zuppann, C. A. (2015). The Quantity-Quality trade-off and the Formation of Cognitive and Non-cognitive Skills. Technical report, National Bureau of Economic Research, WP 21824.

Kim, J. H., Schulz, W., Zimmermann, T., and Hahlweg, K. (2018). Parent–child interactions and child outcomes: Evidence from randomized intervention. *Labour Economics*, 54:152–171.

Lundberg, S. (2005). Sons, daughters, and parental behaviour. *Oxford Review of Economic Policy*, 21(3):340–356.

McHale, S. M., Crouter, A. C., and Whiteman, S. D. (2003). The Family Contexts of Gender Development in Childhood and Adolescence. *Social Development*, 12(1):125–148.

McLanahan, S., Tach, L., and Schneider, D. (2013). The Causal Effects of Father Absence. *Annual Review of Sociology*, 39:399–427.

Mileva-Seitz, V. R., Ghassabian, A., Bakermans-Kranenburg, M. J., van den Brink, J. D., Linting, M., Jaddoe, V. W., Hofman, A., Verhulst, F. C., Tiemeier, H., and van IJzendoorn, M. H. (2015). Are boys more sensitive to sensitivity? parenting and executive function in preschoolers. *Journal of Experimental Child Psychology*, 130:193–208.

Nguyen, G. (2019). Sibling-sex composition, childbearing and female labour market outcomes in Indonesia. *Journal of Population Research*, 36(1):13–34.

Nollenberger, N., Rodríguez-Planas, N., and Sevilla, A. (2016). The Math Gender Gap: The Role of Culture. *American Economic Review*, 106(5):257–61.

Perazzo, I., Salas, G., and Sena, E. (2018). Informe final prácticas de crianza y desarrollo infantil. un estudio longitudinal sobre la bidireccionalidad de este vínculo. Technical report, MIDES, http://dspace.mides.gub.uy:8080/xmlui/handle/123456789/1265

Peter, N., Lundborg, P., Mikkelsen, S., and Webbink, D. (2018). The effect of a sibling's gender on earnings and family formation. *Labour Economics*, 54:61–78.

Pollmann-Schult, M. (2017). Sons, Daughters, and the Parental Division of Paid Work and Housework. *Journal of Family Issues*, 38(1):100–123.

Powell, B. and Steelman, L. C. (1993). The Educational Benefits of Being Spaced Out: Sibship Density and Educational Progress. *American Sociological Review*, pages 367–381.

Rao, N. and Chatterjee, T. (2018). Sibling gender and wage differences. *Applied Economics*, 50(15):1725–1745.

Reynolds, J. R. and Burge, S. W. (2008). Educational expectations and the rise in women's post-secondary attainments. *Social Science Research*, 37(2):485–499.

Reynolds, S. A. (2018). Sex of children and family structure in brazil: father & grandmother bias? *Revista Brasileira de Estudos de População*, 35(1).

Rodríguez-Planas, N. and Nollenberger, N. (2018). Let the girls learn! it is not only about math... it's about gender social norms. *Economics of Education Review*, 62:230–253.

Romano, J. P. and Wolf, M. (2005). Stepwise Multiple Testing as Formalized Data Snooping. *Econometrica*, 73(4):1237–1282.

Romano, J. P. and Wolf, M. (2016). Efficient computation of adjusted p-values for resampling-based stepdown multiple testing. *Statistics & Probability Letters*, 113:38–40.

Rosenzweig, M. R. and Wolpin, K. I. (2000). Natural "Natural Experiments" in Economics. *Journal of Economic Literature*, 38(4):827–874.

Saarela, J. and Finnäs, F. (2014). Sex composition of children, parental separation, and parity progression: Is Finland a Nordic outlier? *Demographic Research*, 30:49–70.

Sayour, N. (2019). The impact of maternal care on child development: Evidence from sibling spillover effects of a parental leave expansion. *Labour Economics*, 58:167–186.

Vladasel, T. (2018). Same, but different?: Birth Order, Family Size, and Sibling Sex Composition Effects in Entrepreneurship. *Swedish Institute for Social Research (SOFI), Working Paper*, 8.

Vogl, T. S. (2013). Marriage Institutions and Sibling Competition: Evidence from South Asia. *The Quarterly Journal of Economics*, 128(3):1017–1072.

2.A Appendix for Chapter 2

Regression analysis: all children with a younger sibling

Table 2.A.1: Summary statistics by sex.

	Girls	Boys
Motor skills		
Gross-motor skills	0.109	-0.086
	(0.654)	(1.018)
Fine-motor skills[a]	0.265	-0.265
	(0.765)	(1.107)
Cognitive skills		
Communication	0.179	-0.124
	(0.742)	(1.144)
Problem-solving	0.200	-0.215
	(0.826)	(1.183)
Non-Cognitive skills		
Socio-personal	0.219	-0.121
	(0.746)	(0.988)
Externalized problems	-0.039	0.152
	(0.914)	(0.967)
Internalized problems	-0.020	0.072
	(0.89)	(0.835)
Total problems	-0.026	0.090
	(0.841)	(0.889)
Observations	224	260

Notes: The table reports means and standard deviations of the ASQ-3 and CBCL tests. The sample consists of all children who have a younger sibling in wave 2.

[a] The sample of fine-motor skills for girls is of 138 and for boys 159.

Table 2.A.2: Summary statistics of all children with a younger sibling. Variables included in the regression analysis.

	Opposite-sex (1)	Same-sex (2)	Difference (3)	t-stat (4)	p-value (5)
Girls					
Age (in months)	51.2	52.1	0.909	0.716	0.475
Race (1=white)	0.78	0.76	-0.021	-0.372	0.710
Household size	5.2	5.0	-0.198	-0.198	0.363
Ln per-capita income	7.8	8.0	0.204	1.463	0.145
Region	0.39	0.48	0.083	1.251	0.211
Mother's age	29.0	29.6	0.605	0.779	0.437
Mother not completed HS	0.60	0.56	-0.038	-0.568	0.570
Mother completed HS	0.26	0.26	0.005	0.089	0.929
Mother completed College	0.15	0.18	0.032	0.684	0.494
First child	0.52	0.67	0.152	2.306	0.021
Second child	0.27	0.15	-0.124	-2.281	0.023
Third born or beyond	0.21	0.18	-0.028	-0.501	0.616
Birth spacing (in months)	31.1	32.2	1.133	0.718	0.474
Sibling's less than 1 year old	0.34	0.32	-0.024	-0.377	0.706
Observations	117	107			
Boys					
Age (in months)	52.9	52.4	-0.489	-0.417	0.677
Race (1=white)	0.81	0.79	-0.015	-0.313	0.755
Household size	5.3	5.4	0.024	0.098	0.922
Ln per-capita income	7.8	7.8	0.034	0.269	0.788
Region	0.44	0.37	-0.070	-1.164	0.244
Mother's age	29.5	29.4	-0.082	-0.100	0.921
Mother not completed HS	0.67	0.66	-0.002	-0.041	0.967
Mother completed HS	0.15	0.16	0.007	0.166	0.868
Mother completed College	0.18	0.18	-0.005	-0.099	0.921
First child	0.63	0.59	-0.032	-0.530	0.596
Second child	0.19	0.24	0.051	0.993	0.321
Third born or beyond	0.18	0.16	-0.019	-0.401	0.688
Birth spacing (in months)	33.9	31.9	-1.993	-1.299	0.195
Sibling's less than 1 year old	0.40	0.35	-0.050	-0.833	0.405
Observations	120	140			

Notes: The table reports means and standard deviations of individual and family characteristics. The sample consists of all children who have a younger sibling in wave 2.

Table 2.A.3: Childhood Development Outcomes (children with a younger sibling)

	ASQ-3					*CBCL*		
	Motor Skills[a]		*Cognitive Skills*[b]			*Non-Cognitive Skills*[c]		
	Gross-motor	Fine-motor	Commu-nication	Problem solving	Socio-personal	Externa-lized	Interna-lized	Total problems
Panel A: Girls								
Opposite-sex mean	0.1451	0.2229	0.1546	0.1237	0.2160	-0.0178	-0.0470	-0.0429
	(0.055)	(0.085)	(0.074)	(0.087)	(0.075)	(0.087)	(0.086)	(0.080)
Panel A.1: No controls								
Same-sex	-0.0527	0.0537	0.0750	0.1682	0.0130	0.0513	-0.0560	-0.0289
	(0.084)	(0.133)	(0.098)	(0.109)	(0.096)	(0.121)	(0.109)	(0.107)
	[0.611]	[0.611]	[0.619]	[0.285]	[0.954]	[0.954]	[0.942]	[0.954]
Prob>F	0.198	0.109	0.004	0.011	0.328	0.430	0.516	0.387
R2	0.016	0.047	0.046	0.071	0.022	0.013	0.009	0.012
Panel A.2: Individual and family characteristics								
Same-sex	-0.062	0.0634	0.0767	0.1562	-0.019	0.005	-0.0983	-0.0705
	(0.080)	(0.129)	(0.097)	(0.103)	(0.097)	(0.118)	(0.106)	(0.105)
	[0.731]	[0.731]	[0.453]	[0.257]	[0.976]	[0.976]	[0.737]	[0.804]
Prob>F	0.017	0.000	0.005	0.000	0.048	0.000	0.000	0.000
R2	0.088	0.198	0.134	0.186	0.096	0.170	0.145	0.172
Observations	224	138	224	224	224	224	224	224
Panel B: Boys								
Opposite-sex mean	0.0154	-0.0763	-0.0869	-0.1764	-0.0486	0.1150	-0.0596	-0.0037
	(0.082)	(0.115)	(0.100)	(0.095)	(0.089)	(0.081)	(0.068)	(0.070)
Panel B.1: No controls								
Same-sex	-0.1815	-0.3317*	-0.0712	-0.0622	-0.1407	-0.0734	-0.2487**	-0.1781*
	(0.127)	(0.174)	(0.141)	(0.147)	(0.125)	(0.119)	(0.099)	(0.107)
	[0.128]	[0.112]	[0.826]	[0.826]	[0.459]	[0.571]	[0.054]	[0.228]
Prob>F	0.214	0.301	0.229	0.228	0.642	0.842	0.094	0.379
R2	0.014	0.024	0.023	0.018	0.007	0.003	0.024	0.011
Panel B.2: Individual and family characteristics								
Same-sex	-0.1776	-0.3160*	-0.0811	-0.0534	-0.1253	-0.0881	-0.2593***	-0.1951*
	(0.127)	(0.169)	(0.141)	(0.140)	(0.124)	(0.117)	(0.097)	(0.104)
	[0.172]	[0.112]	[0.774]	[0.774]	[0.503]	[0.503]	[0.026]	[0.152]
Prob>F	0.175	0.000	0.040	0.005	0.303	0.002	0.000	0.000
R2	0.049	0.213	0.069	0.113	0.052	0.107	0.116	0.118
Observations	260	159	260	260	260	260	260	260

Notes: The table reports the average treatment effect of having a same-sex second-born sibling, for separate regressions of ASQ-3 and CBCL tests, by the sex of the child. Panels A.1 and B.1 do not have controls. Panels A.2 and B.2 include dummies to control for birth order; birth spacing (in months) with the second-born child; a dummy variable equals to one when the sibling is less than one year old; age in months of the child; race (white=1); region (Montevideo=1), the logarithm of people living in the household; the logarithm of per capita income; mother's age; a dummy variable if the mother completed high school, and a dummy variable if the mother is college graduate. Robust standard errors (in parenthesis). Estimates marked *** are significant at 1%, ** at 5%, * at 10%.

[a] Romano-Wolf *p-values* (in square brackets) for motor skills (two coefficients).

[b] Romano-Wolf *p-values* (in square brackets) for cognitive skills (two coefficients).

[c] Romano-Wolf *p-values* (in square brackets) for non-cognitive skills (four coefficients).

Table 2.A.4: Family structure

	Two-parent household	Number of siblings in household	Mother worked
Panel A: Girls			
Opposite-sex mean	0.7692	1.9402	0.5128
	(0.039)	(0.126)	(0.046)
Panel A.1: No controls			
Same-sex	-0.0373	0.0520	0.0794
	(0.061)	(0.102)	(0.068)
Prob>F	0.909	0.000	0.158
R2	0.003	0.717	0.022
Panel A.2: Individual and family characteristics			
Same-sex	-0.0619	0.0629	0.0268
	(0.055)	(0.086)	(0.056)
Prob>F	0.000	0.000	0.000
R2	0.233	0.796	0.362
Observations	224	224	224
Panel B: Boys			
Opposite-sex mean	0.7750	1.8333	0.5333
	(0.038)	(0.137)	(0.046)
Panel B.1: No controls			
Same-sex	-0.1207**	0.05240	0.0118
	(0.055)	(0.110)	(0.061)
Prob>F	0.109	0.000	0.532
R2	0.0214	0.702	0.009
Panel B.2: Individual and family characteristics			
Same-sex	-0.1240***	0.04032	-0.0045
	(0.048)	(0.087)	(0.052)
Prob>F	0.000	0.000	0.000
R2	0.2201	0.783	0.315
Observations	260	260	260

Notes: The table reports the average treatment effect of having a same-sex second-born sibling, for separate regressions of family characteristics, by the sex of the child. The effects relating to the binary variables have been estimated using a logit regression model. The reported coefficients are marginal effects evaluated at the average probability and then averaged over the sample. Panels A.1 and B.1 control for birth order. Panels A.2 and B.2 include dummies to control for birth order; birth spacing (in months) with the next youngest sibling; a dummy variable equals to one if the sibling is less than one year old, age in months of the child; race (white=1); region (Montevideo=1), logarithm of people living in the household; the logarithm of per capita income; mother's age; a dummy variable if the mother completed high school and a dummy variable if the mother is college graduate. Standard errors (in parenthesis) are adjusted for clustering at the family level. Estimates marked *** are significant at 1%, ** at 5%, * at 10%.

Table 2.A.5: Parental investment and expectations

	Parental investment			*Health-care*
	Active Time	Attending School	College Expectations	Flu immunization
Panel A: Girls				
Opposite-sex mean	0.0610	0.8120	0.6068	0.2308
	(0.085)	(0.036)	(0.036)	(0.039)
Panel A.1: No controls				
Same-sex	0.1664	0.0607	0.0607	0.0461
	(0.119)	(0.045)	(0.045)	(0.058)
Prob>F	0.309	0.003	0.003	0.066
R2	0.021	0.091	0.091	0.030
Panel A.2: Individual and family characteristics				
Same-sex	0.1626	0.0687	0.0687	0.0063
	(0.117)	(0.042)	(0.042)	(0.056)
Prob>F	0.071	0.000	0.000	0.000
R2	0.104	0.235	0.235	0.166
Observations	224	224	224	224
Panel B: Boys				
Opposite-sex mean	0.0359	0.9000	0.9000	0.2417
	(0.086)	(0.027)	(0.027)	(0.039)
Panel B.1: No controls				
Same-sex	-0.3713***	-0.0449	-0.0449	0.0041
	(0.138)	(0.039)	(0.039)	(0.053)
Prob>F	0.032	0.039	0.039	0.054
R2	0.035	0.003	0.003	0.690
Panel B.2: Individual and family characteristics				
Same-sex	-0.3759***	-0.0603	-0.0603	0.0224
	(0.138)	(0.041)	(0.041)	(0.048)
Prob>F	0.003	0.042	0.042	0.049
R2	0.113	0.000	0.000	0.000
Observations	260	260	260	260

Notes: Standard errors (in parenthesis) are adjusted for clustering at the family level. Estimates marked *** are significant at 1%, ** at 5%, * at 10%. Other notes see Table 2.A.4.

Table 2.A.6: Parental practices

	Parenting styles		*Home environment*	
	Authoritarian	Authoritative	Warm	Harsh
Panel A: Girls				
Opposite-sex means	-0.1085	0.0144	1.5214	0.5983
	(0.090)	(0.073)	(0.128)	(0.099)
Panel A.1: No controls				
Same-sex	0.0327	-0.1825	0.2211	0.0200
	(0.129)	(0.132)	(0.190)	(0.129)
Prob>F	0.223	0.361	0.022	0.122
R2	0.023	0.018	0.044	0.026
Panel A.2: Individual and family characteristics				
Same-sex	0.0714	-0.2029	0.2679	0.0293
	(0.131)	(0.128)	(0.189)	(0.124)
Prob>F	0.072	0.019	0.000	0.007
R2	0.086	0.103	0.162	0.135
Observations	224	224	224	224
Panel B: Boys				
Opposite-sex mean	0.0919	0.0652	1.8500	0.8917
	(0.100)	(0.076)	(0.135)	(0.111)
Panel B.1: No controls				
Same-sex	0.0086	0.0631	0.0018	-0.0239
	(0.141)	(0.101)	(0.190)	(0.164)
Prob>F	0.029	0.462	0.712	0.712
R2	0.043	0.010	0.005	0.006
Panel B.2: Individual and family characteristics				
Same-sex	-0.000	0.0760	0.0180	-0.0385
	(0.132)	(0.102)	(0.180)	(0.162)
Prob>F	0.001	0.241	0.001	0.075
R2	0.137	0.059	0.124	0.078
Observations	260	260	260	260

Notes: Standard errors (in parenthesis) are adjusted for clustering at the family level. Estimates marked *** are significant at 1%, ** at 5%, * at 10%. Other notes see Table 2.A.4.

Chapter 3

Collective Labor Supply, Divisions of Domestic Work and Intra-household Bargaining

3.1 Introduction

The gender division of labor has changed since the Second World War. Women's labor supply has drastically increased, and men's has decreased at a slower and more stable rate. On the other hand, the household division of work has become less gender specialized. This latter result is explained by the interaction of different patterns regarding the increase in the educational level of women, changes in fertility, marriage, divorce patterns, and by technological advances that have allowed the goods produced in the home to be marketable (Lundberg, 2010; Greenwood et al., 2016). However, gender differences in the division of domestic work persist, with women continuing to perform most of the household chores.

The intra-household division of labor may be related to the relative resources of each partner. However, it may also reflect internalized gender norms, which define what is considered acceptable behavior for men and women (Agarwal, 1997; Pearse and Connell, 2016). Households with an egalitarian or non-traditional division of domestic labor are more likely if partners have similar economic resources or the influence of gender norms is low (Seiz, 2021).

The empirical evidence regarding the relationship between intra-household bargaining and gender division of domestic work is scarce in the economic literature and focuses on high-income countries. This paper contributes to the literature by providing evidence from

a developing country on the intra-household bargaining of heterosexual couples with different gender norm attitudes towards the division of domestic work and its impact on the couple's labor supply. In particular, the aim of this paper is to compare the decision-making process in families with traditional, egalitarian, and non-traditional gender role attitudes.

The model of labor supply with distribution factors proposed in Chiappori et al. (2002) is used to identify the derivatives of the sharing rule for each household type. Two distribution factors are used in this paper: the non-labor income difference between partners and the condition of being married versus cohabiting. According to the collective model, the non-labor income difference affects the decision-making process by giving more power to the richer partner, which translates into a decreased pattern of labor supply for him/her. This variable is a direct test of the validity of the unitary model, which imposes the *income-pooling hypothesis*. This hypothesis states that the source of the income is not relevant to the final allocation of outcomes, but the total income is. The rejection of the income-pooling hypothesis implies that the unitary model is not suitable for use when analyzing the household decision process (Browning et al., 1994).

Most studies that use collective models to analyze the labor supply decisions suppose that individuals are single or married, leaving aside cohabitation decisions. When this decision is considered, results compare labor supply responses of those who choose to be married and those who cohabit. In practice, the most significant difference in choosing marriage versus cohabitation is that the costs of household formation and dissolution are higher for married couples.

Prenuptial contracts, which define in the case of divorce how property rights are distributed between the couple, can reduce the costs of divorce (Bayot and Voena, 2015). Therefore, if no prenuptial contracts are established, marriage can be thought of as a distribution factor that protects women who specialize in domestic work from divorce (Gemici and Laufer, 2011). If this mechanism holds, the bargaining power of married women should be higher than those who are cohabiting, and the corresponding negative (positive) correlation between female (male) labor supply and being married should be observed.

This paper relates to the literature that analyzes the interactions of social norms on a variety of labor market outcomes. Fernández (2013) has developed a learning model to show how the probability of being a working mother generates an information update on the value of labor, changing the female labor market participation in the long term. The question "Do you approve or disapprove of a married woman earning money in business or industry if she has a husband capable of supporting her?" is used to compare the pre-

dictions of the model to changes in social attitudes. Bertrand et al. (2015) analyzed the correlation of social norms on women's labor force participation and gender gap in incomes using the question "If a woman earns more money than her husband, it's almost certain to cause problems"; using the same definition, Galván (2021) analyzed how gender role attitudes in Uruguay affect women's and men's probability of employment in a formal job. Goussé et al. (2017) constructed an index of family values to account for marriage decisions and intra-household bargaining regarding consumption and labor supply. More recently, Bertrand et al. (2021) have studied the interactions between economic opportunities, gender norms and marriage rates among skills groups to account for the diverging patterns in labor supply participation of married women in a pool of developed countries. To measure gender norms they use two questions: "When jobs are scarce, men have more right to a job than women" (from the World Value Survey) and "A man's job is to earn money; a woman's job is to look after the home and family" (from the International Social Science Program). However, none of these studies have directly analyzed the correlation between social attitudes towards the division of domestic work and the intra-household bargaining process. I measure these social norms in Uruguay using the question "Is she/he the one who mainly does the household chores?".[1]

Uruguay has a long history in terms of women's rights, being the first country in Latin America to grant a divorce by the sole will of the woman throughout the territory in 1913. It was also the first to grant women fully equal civil and universal suffrage in the constitutional reform of 1917, though this suffrage wasn't exercised until the national general elections of 1938. In recent years, there have been several legislative advances such as the gender-based violence law (2004), cohabitation union law (2006), quota law (2009), gender identity law (2009), voluntary interruption of pregnancy law (2012), and parental leave law (2013), among others.

In Uruguay, the average level of literacy, life expectancy, and gross per capita income is above the median of other Latin American countries. It has a high level of human development (Human Development Index of 0.817, 2017) and has been considered a high-income country according to the World Bank since 2013 (PNUD, 2020). However, there are still significant gaps in the Gender Inequality Index (GII) with Uruguay's value of 0.288 placing it in 62nd place out of 148 countries (PNUD, 2020). According to Batthyány et al. (2015), the challenges in Uruguay to close the remaining gaps are related to promoting policies that

[1]Uruguay's *"Encuesta Continua de Hogares"* survey asks this question to all household members over 14 years of age.

guarantee women's access to the labor market, stable and quality jobs, and policies that encourage cultural change within households in the distribution of domestic work between men and women.

In this context, we would expect that households with traditional gender role attitudes conform more with the "breadwinner" hypothesis, which gives more power to men, and therefore, the power of women in the negotiation process is lower than in egalitarian or non-traditional households. Furthermore, according to the collective model of labor supply, a shift in the bargaining power of women should increase the domestic work of both spouses.

This paper contributes to the literature on gender role attitudes by analyzing the decision-making process within families with different social norms. In particular, it provides new evidence on the relationship between labor supply choices and the bargaining power within households, the difference between partners' non-labor income and whether or not they are married. It also supports collective rationality in a non-developed country and rejects the "income pooling" hypothesis. This has a direct impact on policy design, since shifts in distribution factors such as a conditional cash transfer, will affect the bargaining power of household members, leading to changes in the labor supply choices of each household member.

The layout of the paper is as follows. Section 3.2 presents the collective model of labor supply with distribution factors, the collective rationality tests, and the derivatives of the sharing rule. Section 3.3 describes the empirical approach, the data and the variable that approximates the measure of social attitudes regarding the division of domestic work, as well as the descriptive statistics. Section 3.4 discusses the results of the collective model of labor supply among household types and presents a sensitivity analysis of selection into employment. Finally, Section 3.5 concludes.

3.2 Theoretical framework

This section presents the basic model of collective labor supply with distribution factors as described by Chiappori et al. (2002), which is applied to households with traditional, egalitarian and non-traditional social norms. The household is composed of two individuals with different preferences and utility functions, which in general, are assumed to be altruistic. These agents know each other's preferences, interacting in a stable decision process, which leads to Pareto-efficient outcomes. The individuals of working age have a general

utility function that depends on consumption and leisure of both partners:

$$U^i = U^i(1 - h^f, C^f, 1 - h^m, C^m, \mathbf{z}), \quad i = f, m \tag{3.1}$$

where f=female and m=male, U^i are strictly quasi-concave functions, increasing and continuously differentiable, h^i is the member i's labor supply (with $0 \leq h^i \leq 1$), C^i is the member i's consumption of a Hicksian composite good whose price has been normalized to one, and $\mathbf{z}$ is a vector of preference variables, which include individual and family characteristics. The household budget constraint is:

$$C^f + C^m \leq w_f h^f + w_m h^m + y \tag{3.2}$$

where w_i is the wage rate of individual i, and y is the household non-labor income. Under the collective model, the decision process is Pareto efficient, which implies that for any given set of $(w_f, w_m, y, \mathbf{z}, \mathbf{s})$, there exists a weighting factor $\mu(w_f, w_m, y, \mathbf{z}, \mathbf{s}) \in [0, 1]$, that solves the following maximization program:

$$\begin{aligned} \max_{\{h^f, h^m, C^f, C^m\}} \; & \mu(w_f, w_m, y, \mathbf{z}, \mathbf{s}) U^f + (1 - \mu(w_f, w_m, y, \mathbf{z}, \mathbf{s})) U^m \\ \text{s.t.} \; & C^f + C^m \leq w_f h^f + w_m h^m + y \\ & 0 \leq h^i \leq 1, \quad i = f, m \end{aligned} \tag{P1}$$

where $\mathbf{s}$ is a vector of distribution factors -i.e., variables that affect the bargaining position within the household- that does not affect preferences or the budgetary restriction of the household (Browning et al., 2014). As the solution is Pareto-efficient, an increase in μ reflects a movement along the Pareto frontier, in the same direction that gives more utility to the female member of the household. Thus the Pareto weight can be considered a direct interpretation of power within the family. At the same time, these weights do not imply that both members have to agree on every decision. Pareto weights ensure that whatever the final solution is, no resource will be unused. The final solution in the Pareto frontier depends on the value of the parameters involved: the individual i's wage $(w_i, i = f, m)$, the non-labor income of the family (y), the individual and family characteristics $(\mathbf{z})$ and the distribution factors $(\mathbf{s})$, and therefore, the Pareto weight.

When preferences are egoistic each partner takes their own preferences into account.[2]

[2]The model allows extensions for caring preferences and public goods, e. g., Blundell et al. (2005).

The utility function of member $i = f, m$ is $U^i(1 - h^i, C^i, \mathbf{z})$, where U^i is strictly quasi-concave, increasing and continuously differentiable. Chiappori (1992) shows that the solution to (P1) is equivalent to the following decentralized program[3]:

$$\begin{aligned} \max_{\{h^i, C^i\}} \; & U^i(1 - h^i, C^i, \mathbf{z}) \\ \text{s.t.} \; & C^i \leq \phi^i + w_i h^i \\ & 0 \leq h^i \leq 1 \end{aligned} \tag{P2}$$

where $\phi^f = \phi$ represents the female's share of non-labor income y, while the male's share is given by $\phi^m = y - \phi$. The equivalence between program (P1) and program (P2) is provided by the second fundamental welfare theorem, so that for any solution $(w_f, w_m, y, \mathbf{z}, \mathbf{s})$ of program (P1), there exists a *sharing rule* (ϕ) for which $(w_f, w_m, y, \mathbf{z}, \mathbf{s})$ is also the solution of the program (P2). According to (P2), the decision process is done in two steps: in the first, members of the family decide how much of the non-labor income is allocated to each member, and in the second, each individual chooses their own labor supply and private consumption. The sharing rule can be positive or negative because it can include the allocation of both labor and non-labor income between partners.

The solution to program (P1) yields the following set of Marshallian labor supply equations:

$$\begin{aligned} h^f &= h^f(w_f, w_m, y, \mathbf{s}, \mathbf{z}) \\ h^m &= h^m(w_f, w_m, y, \mathbf{s}, \mathbf{z}) \end{aligned}$$

Whereas, the Pareto-efficient decision to (P2) gives the following set of Marshallian labor supply functions:

$$h^f = H^f(w_f, \phi(w_f, w_m, y, \mathbf{s}, \mathbf{z}), \mathbf{z}) \tag{3.3}$$

$$h^m = H^m(w_m, y - \phi(w_f, w_m, y, \mathbf{s}, \mathbf{z}), \mathbf{z}) \tag{3.4}$$

The results in equations (3.3) and (3.4) show that only own wage matters when determining the individual's demand for leisure. The wage rate of the partner, the non-labor income (y), and distribution factors $(\mathbf{s})$ affect the demand for leisure only through the sharing rule, based on an income effect, being leisure a normal good. Therefore, factors that improve a

[3]Proposition 2 in Chiappori et al. (2002).

woman's bargaining power reduce her labor supply and increase her partner's supply.

When the specified model uses at least two distribution factors, collective rationality can be empirically tested. Browning et al. (1994) and Bourguignon et al. (2009) show that the necessary and sufficient conditions for a collective model are:

$$\frac{\partial h^f / \partial s_\ell}{\partial h^f / \partial s_1} = \frac{\partial h^m / \partial s_\ell}{\partial h^m / \partial s_1}, \quad \text{for } \ell = 1, ..., L \tag{3.5}$$

where s_ℓ is a distribution factor. Equation (3.5) told us that the marginal rate of substitution between labor supply and any given pair of distribution factors has to be proportional between the two members of the couple.[4] The proportionality test holds if the distribution factors only affect the decision process through the one-dimensional function ϕ. If conditions in equation (3.5) are not rejected, the allocation of resources between partners is consistent with a static definition of efficiency (Chiappori and Mazzocco, 2017).[5] There is another test for collective rationality which imposes restrictions on the Slutsky matrix, but it is only testable when panel data with price variation is available (Browning et al., 2014).

Chiappori (1988, 1992) and Chiappori et al. (2002) have proved the conditions to recover the sharing rule up to an additive constant, from the observation of individual labor supply as a function of wages, non-labor income and distribution factors. Under the assumption that the sharing rule is increasing in non-labor income, and if the partial derivative of the sharing rule to the distribution factors are different from zero $\left(\partial\phi/\partial s_\ell \neq 0, \text{ i.e., } \frac{\partial h^f/\partial s_\ell}{\partial h^f/\partial y} \neq \frac{\partial h^m/\partial s_\ell}{\partial h^m/\partial y}, \text{ Browning et al. (2014)}\right)$, the following relationships holds:[6]

$$\phi_{w_f} = \frac{\frac{h^m_{w_f}}{h^m_y} \cdot \frac{h^f_{s_\ell}}{h^f_y}}{\frac{h^m_{s_\ell}}{h^m_y} - \frac{h^f_{s_\ell}}{h^f_y}}, \quad \phi_{w_m} = \frac{\frac{h^f_{w_m}}{h^e_y} \cdot \frac{h^m_{s_\ell}}{h^m_y}}{\frac{h^m_{s_\ell}}{h^m_y} - \frac{h^f_{s_\ell}}{h^f_y}}, \quad \phi_y = \frac{\frac{h^m_{s_\ell}}{h^m_y}}{\frac{h^m_{s_\ell}}{h^m_y} - \frac{h^f_{s_\ell}}{h^f_y}}, \quad \phi_{s_\ell} = \frac{\frac{h^f_{s_\ell}}{h^f_y} \cdot \frac{h^m_{s_\ell}}{h^m_y}}{\frac{h^m_{s_\ell}}{h^m_y} - \frac{h^f_{s_\ell}}{h^f_y}}, \quad \forall \ell = 1, ..., L \tag{3.6}$$

Then, for a given empirical specification of the labor supply functions, we can recover

[4] Proposition 1 in Chiappori et al. (2002).

[5] This is equivalent to saying that it is a test for Pareto efficiency, the main assumption in the collective model of labor supply.

[6] Where $h^i_j = \frac{\partial H^i}{\partial \phi}\frac{\partial \phi}{\partial j}$, with $i = f, m$ and $j = w_f, w_m, y, s_\ell$.

the sharing rule as an additive function of these partial derivatives.

In the empirical literature, different proxies have been used for distribution factors: local sex ratios, divorce laws, abortion legalization, relative age, relative non-labor income, married (vs. cohabiting), relative incomes, relative wages, relative education, background family factors, control of land, previous children, reported influence within the household, single parents benefits, etc. (Chiappori et al., 2002; Oreffice, 2007; Rapoport et al., 2011; Haddad, 2015; Oreffice, 2011). Some of these distribution factors will be used in the following sections to analyze the bargaining power within households with different divisions of domestic work.

The division of domestic and market work may in part be linked to the relative resources of each partner, although it may also reflect internalized gender norms. Furthermore, different specialization patterns within the family affect the outside options, which, in turn, affect the negotiation process. Analyzing the bargaining process in households with traditional, egalitarian, and non-traditional gender norms will give some insights into the balance of power within each household type.

3.3 Empirical specification, data and social attitudes

3.3.1 Empirical specification

To estimate the collective model it is assumed that the system of labor supply equations of the household members has the following semi-logarithmic form:

$$h^f = f_0 + f_1 \log w_f + f_2 \log w_m + f_3 y + f_4 s_1 + f_5 s_2 + \mathbf{f_6'z} + \epsilon_f \tag{3.7}$$

$$h^m = m_0 + m_1 \log w_f + m_2 \log w_m + m_3 y + m_4 s_1 + m_5 s_2 + \mathbf{m_6'z} + \epsilon_m \tag{3.8}$$

where h^i is the hours usually worked in a week, w_i is the hourly wage of partner i, y is the total non-labor income of the household in thousands, s_1 and s_2 are two distribution factors, $\mathbf{z}$ are individual and family characteristics (preferences), and ϵ_f and ϵ_m are the error terms which may or may not be correlated. The first distribution factor (s_1) is the difference between non-labor income assignable to the male partner and the non-labor income of the female partner measured in thousands. The second (s_2) is the condition of being married, or not.[7]

[7]Following Browning et al. (1994) and Oreffice (2011), I use the difference between non-labor income of men and women to avoid having missing values in the ratio.

According to the collective model of labor supply, the member of the family that faces a positive difference in non-labor income has more bargaining power within the household, decreasing their own-labor supply and increasing the labor supply of their partner. Thus, if males earn more non-labor income than females, the coefficient f_4 should be positive, and m_4 should be negative.

Regarding being married versus cohabiting, the bargaining power of partners can be associated with the community property regime. Bayot and Voena (2015) analyze changes in labor force participation, wages, and specialization patterns when there are prenuptial agreements in place for married couples. They found that women engaged in marriages with community property contracts participate less in the labor market, have lower wages, and the likelihood of specialization is higher than for women in a separate community regime. They suggest that community property contracts provide insurance against divorce for wives who self-select into a traditional division of work. In such a regime, females decrease their labor supply while males increase theirs. Therefore, if this mechanism holds, the coefficient f_5 should be negative while m_5 should be positive.

While no differences are expected in the sign of these distribution factors concerning social norms on the division of domestic work, differences in magnitude may be observed due to the differences in semi specialization patterns. Specifically, traditional households are associated with the "breadwinner" norm, which results in men having a higher position in the negotiation process.

In order to test the validity of the collective model, the following necessary and sufficient condition must hold:

$$\frac{m_4}{f_4} = \frac{m_5}{f_5} \tag{3.9}$$

Given $\frac{\partial h^f/\partial s_\ell}{\partial h^f/\partial y} \neq \frac{\partial h^m/\partial s_\ell}{\partial h^m/\partial y}$ for at least one ℓ, the derivatives of the sharing rule are:

$$\phi_{w_f} = \frac{1}{\Delta}\frac{m_1 f_4}{w_f}, \quad \phi_{w_m} = \frac{1}{\Delta}\frac{f_2 m_4}{w_m}, \quad \phi_y = \frac{f_3 m_4}{\Delta}, \quad \phi_{s_\ell} = \frac{f_\ell m_4}{\Delta} \tag{3.10}$$

where $\Delta = f_3 m_4 - f_4 m_3$. To recover the sharing rule, up to an additive constant $k(z)$, we have to solve the system of differential equations in (3.10), resulting in:

$$\phi = \frac{1}{\Delta}\left(m_1 f_4 \log w_f + f_2 m_4 \log w_m + f_3 m_4 y + f_4 m_4 s_1 + f_5 m_4 s_2\right) + k(z) \tag{3.11}$$

The conditions imposed by the theoretical model will now be tested and presented by esti-

mating the labor supply equations of the family.[8]

3.3.2 Data and social attitudes towards the division of domestic work

This paper uses the Household Survey of Uruguay (*Encuesta Continua de Hogares*, ECH)), a survey undertaken annually by the National Statistics Institute (*Instituto Nacional de Estadística*, INE).[9] The survey collects information about socioeconomic characteristics of the household, including the usual hours of labor in principal and secondary occupations, the wage rate in each type of occupation, the sector of activity, and non-labor incomes for each partner, as well as a variable that reveals which members of the family do what housework. The main sample consists of heterosexual working couples (married or cohabiting) with both partners between 23 and 64 years old, excluding extended families. Self employed individuals were excluded from the sample so that non-labor income corresponds to non-earned income (Vermeulen, 2005; Oreffice, 2011). Additionally, the sample is restricted to couples for which information is complete for both members and for which either the male of female partner self-reported the information.

The ECH reports for each person over 14 years of age in the household "Is (name) the one who *mainly* does the housework?" and the possible answers are "Yes" and "No". Using this variable, the sample is divided into families attached to traditional gender roles where only women do domestic tasks, egalitarian households where both members of the couple perform housework, and non-traditional families where the women do not perform housework. The evolution of the self-reported division of domestic work within households in Uruguay from 1991 to 2019, in the same restricted sample as used in this paper, is presented in Figure 3.1a. The discontinuity in 2010 reflects a change in the phrasing of the question.[10] According to the data, the percentage of households following a traditional division of labor has decreased sharply since the beginning of the nineties, while egalitarian households have emerged in their place. The stability in the proportion of households self-reported as non-traditional is noteworthy, remaining at around 5% throughout the period. The over-

[8]The sharing rule cannot be recovered because the vector given by $\mathbf{z}$ affects both preferences and the sharing rule. If we want to recover the sharing rule, we have to make additional assumptions (for example that there are not greater differences between singles and married individuals).

[9]In this study, I use cross-sectional data from the *Encuesta Continua de Hogares* from 1991 to 2019 to analyze the trends concerning the division of domestic work within households and variables related to the labor market, and data of the *Encuesta Continua de Hogares* from the year 2019 to analyse the intra-household bargaining of heterosexual couples with different gender norm attitudes towards the division of domestic work.

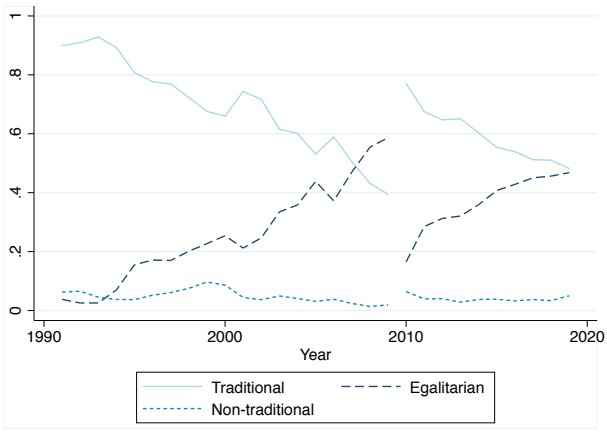
[10]Until 2010, the ECH asked each person over 14 years of age in the household "Is (name) the one who does the housework?", without the adverb "*mainly*", and the possible answers were "Yes" and "No".

all picture may reflect changes in internalized gender norms or at least the perception of them within the households. Similarly to in the work of Fernández (2013), Bertrand et al. (2015), Goussé et al. (2017), Bertrand et al. (2021) and Galván (2021), this paper will use the self-reported perception of division of domestic work to approximate social norms within households.

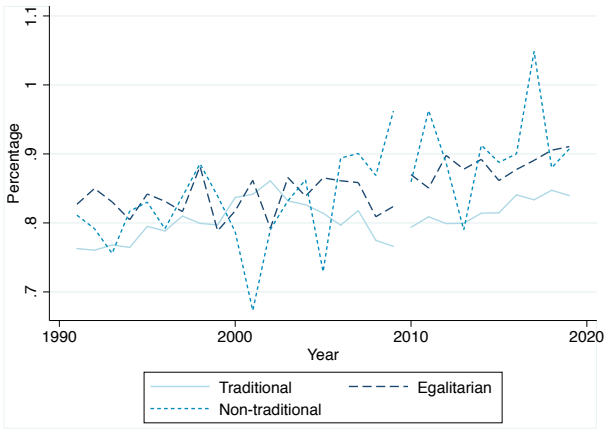
To give some insights into the possible correlation between social norms regarding the division of domestic work and the variables related to the labor market, Figure 3.1 includes the evolution of the raw wage gap between men and women (panel 3.1b), and the hours of work in the market by gender (panels 3.1c and 3.1d). The average wage gap (mean wage of females over mean wage of males) narrows from 0.77 to 0.87 between 1991 and 2019, decreasing from 0.76 to 0.84 in traditional homes, while in egalitarian households this gap narrows from 0.83 to 0.91, and in non-traditional couples from 0.81 to 0.91.

Figures 3.1c and 3.1d presents the evolution of usual hours of market work in Uruguay from 1991 to 2019, for females and males, respectively. The average hours of market work of females in traditional homes decreased from 37 in 1991 to 35 in 2019, whereas, in egalitarian households it remained stable at 38 hours throughout the period. The mean hours of work of women in non-traditional homes increased from 38 to more than 40 hours of work per week. Since early 2000, this difference has widened, driving females in non-traditional and egalitarian households to work more than females in traditional ones. Males, on the contrary, show a downward tendency in the average hours of market work in all household types, but most remarkably in egalitarian and non-traditional homes, decreasing from 52 hours at the beginning of the nineties to 45 hours in 2019. Therefore, the gender gap in hours of market work has narrowed faster in egalitarian and non-traditional homes than in traditional households.

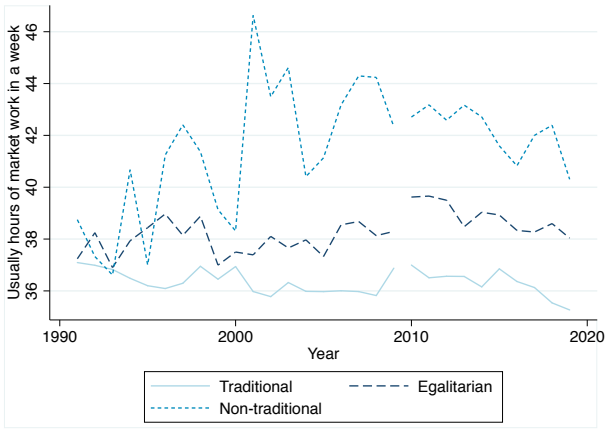
In summary, there seems to be a positive correlation between the evolution of the division of domestic work towards a more egalitarian distribution and variables related to the labor market.



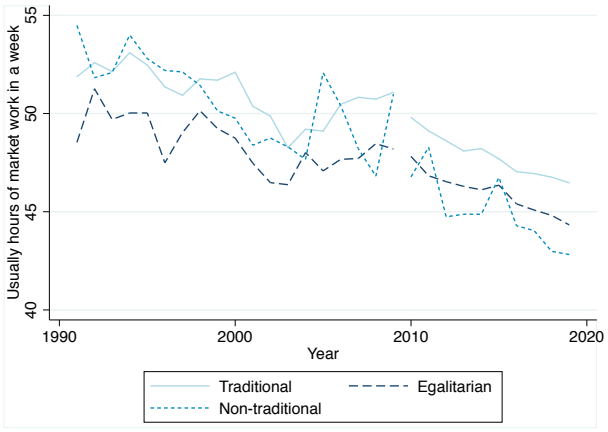
(a) Division of domestic work



(b) Raw wage gap between men and women



(c) Hours of market work female partner



(d) Hours of market work male partner

Figure 3.1: Division of domestic work, raw wage gap and usual hours of work in the labor market in Uruguay, 1991-2019.

Source: Author's calculations based on *Encuesta Continua de Hogares* from 1991 to 2019.
Note: Heterosexual working couples between 23 and 64 years of age.

According to the Time Use Survey of Uruguay from 2013, *Encuesta de Uso del Tiempo (EUT, 2013)*, in households classified in this paper as traditional, 23% of the total hours of domestic work were performed by the male partner, while this share increased to 36% in egalitarian and to 42% in non-traditional households.[11] This pattern is reflected in Figure 3.2, where the distribution of hours of housework between partners is presented through kernel density functions of the share of hours of domestic work done by males for tra-

[11]The EUT (2013) is representative of the Uruguayan population and is a subsample of the ECH survey. It also asks each person in the household about housework.

ditional, egalitarian, and non-traditional households, each shown separately. Most traditional homes are in the lowest share levels, while egalitarian and non-traditional homes are situated to the right. In sum, women do most of the domestic work and the share of hours of housework of males increases from traditional, through egalitarian to non traditional homes. Although the correlation is not perfect, I will use this proxy of the division of domestic work to approximate households with different gender norms. In general, the definition moves from highly specialized households to less specialized: "traditional", "egalitarian", and "non-traditional".

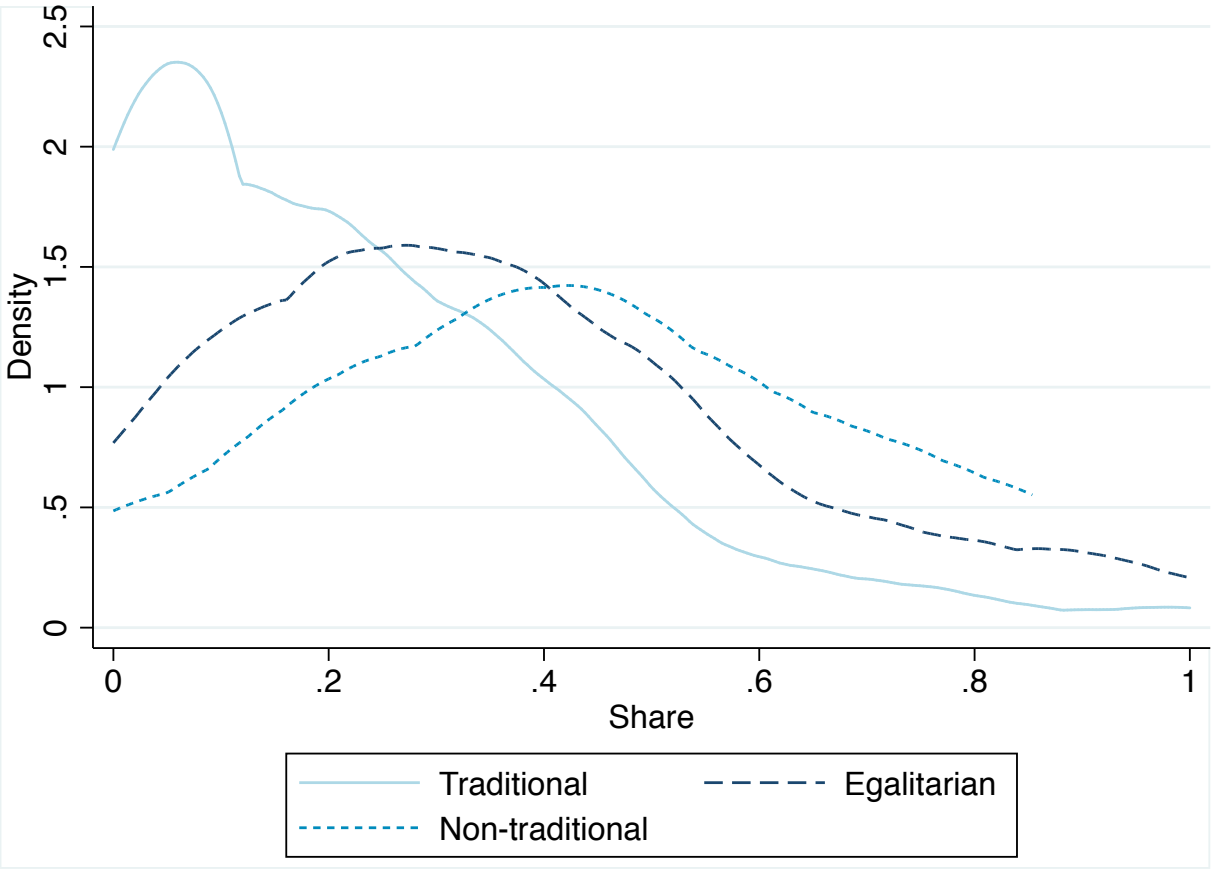


Figure 3.2: Share of male hours of housework on total hours of housework, EUT (2013).

Source: Author's calculations based on *Encuesta de Uso del Tiempo*, 2013.
Note: Heterosexual working couples between 23 and 64 years of age.

3.3.3 Descriptive statistics

Table 3.1 presents the descriptive statistics for the full sample, composed of a total of 4,276 heterosexual couples, as well as the descriptive statistics for the sample split according to the division of domestic work between females and males within the household. The dependent variable is the usual weekly hours of work in the labor market spent on all occupations for each member of the couple. Females work on average 36.3 hours per week, while males work 45.8. The number of hours of market work for females in households that are more attached to traditional gender roles is lower than in egalitarian and non-traditional homes. Among non-traditional couples, the mean hours of market work for females is 40.2 while males work 41.7, in egalitarian couples females work on average 37.7 hours and males 43.9, while in traditional households females work 34.6 and males almost 46 hours per week.

The wage variable includes all individual earnings related to work (in cash or in-kind, valued on the information provided by the survey), expressed all in US dollars as of 2019. In all types of households, the hourly wage of men is higher than that of women, but it narrows when we move from couples with traditional gender roles to those with a non-traditional division of domestic work. The non-labor income variable includes profits, interests, rents, and non-labor income assignable to each member of the couple, net of transfers to other households. The average non-labor income is $64 per month, and it is higher in households with a non-traditional distribution of domestic work.

The non-labor income assignable to each member of the couple is computed as the difference between that of the man and that of the woman. This variable includes capital gains, transfers from contributory programs (e.g., pensions, unemployment insurance, maternity leave), noncontributory pensions, voluntary transfers from other households (grants, subsidies, or donations), family allowances (conditional cash transfers-CCT), and food transfers. CCT and food transfers are means-tested programs. The CCT program consists of a cash transfer targeted at households with children younger than eighteen years of age and is conditional on school attendance and health assessments. The food transfer program, *Tarjeta Uruguay Social*, is a transfer made through a debit card that allows households to purchase food and personal hygiene products. These two programs are targeted at vulnerable homes.

The overall non-labor income difference between males and females is negative and calculated to be $3 per month. In traditional and egalitarian homes, women have on average higher non-labor incomes than men. Conversely, in homes with a non-traditional gender division of domestic work, the mean difference is positive, at approximately $25. The non-

labor income difference is directly related to the *income-pooling hypothesis*, which states that only household income matters for the final allocation of outcomes and not the source of the income (Browning et al., 2014). The rejection of this hypothesis implies that the unitary model is not suitable for use when analyzing the household decision process (Browning et al., 1994).

Married couples make up 51% of the sample while the remaining 49% cohabit. The marriage rate is similar between traditional and non-traditional homes, whereas it decreases in egalitarian households to 48%. Cabella and Fernández Soto (2017) analyze changes in the profile of young cohabitants in Uruguay between 1990 and 2015. They show that this type of union is more common amongst young people, its prevalence decreases with age and there are no differences among educational groups. In 1990 the proportion of individuals aged 20 to 24 living in consensual unions accounted for 20% of the total number of young people in a partnership, whereas in 2000 this proportion rose to almost 50%, and reached 90% in 2015. For individuals aged 40 to 44, these consensual unions increased from 10% in 1990 to approximately 40% in 2015 (Cabella and Fernández Soto, 2017).

A law of cohabitation unions was enacted in 2006. This law ensures that spouses in cohabitation unions have the same rights as married couples. To form a legal cohabitation union it is necessary to prove five years of prior cohabitation, and follow legal procedures. In Uruguay, the default regime of property at the time of marriage or legal cohabitation is community property rights. Before marriage or legal cohabitation, the division of joint property can be established. According to a financial survey carried out in 2016 in Uruguay (*Encuesta Fianciera de los Hogares Uruguayos*, EFHU-2), only 10.4% of individuals who are in a union have separate contracts rights, with rates of 12% among married people and 3.6% among cohabitants. The proportion of separate contract rights rose to 23.5% and 15.8% among individuals divorced from a previous marriage or separated from cohabitation.[12] In this setting, marriage can be a proxy for insurance against divorce and thus may play a significant role as a distribution factor.

Regarding socioeconomic characteristics, the mean age of women in the sample is 40.3, while it is 42.7 for men. Of all respondents, 87% of them report themselves to be white. The average number of years of education is 12.0 and 10.8, for females and males, respectively. As in Goussé et al. (2017), the sample shows that more educated, younger, and richer individuals tend to have a more egalitarian division of domestic work, i.e., are less attached

[12]The EFHU-2 does not have information on the labor market outcomes of the couples. For this reason, this survey is not suitable for use in the analysis of premarital contracts as a distribution factor.

to gender roles. The mean number of children the couple has together is 0.82, while the mean number of children belonging to only the female partner is 0.13, and to only the male partner 0.02. Overall, females own the home in 8.3% of the sample, whereas males own it in 6.4% of cases. This difference widens, in favor of women, in non-traditional homes in which women own the home in 11.4% of the households and men 7.3%. Approximately 61% of the sample is located in the capital city or its metropolitan area, while almost 15% is from the North, and the remaining 25% is from the South. Traditional gender roles prevail in the South and North regions. Finally, the data reports which partner responded to the survey. Since this variable biases the estimates, it was introduced as an additional control.[13] In the whole sample, the female partner responded to the survey in 67% of cases. This rate decreased to 51% in non-traditional and 61% in egalitarian homes, while it reached almost 74% in traditional households.

[13] Amábile et al. (2021) show the importance of accounting for respondent gender when only one individual reports the information from household members.

Table 3.1: Summary statistics.

	All couples		Traditional		Egalitarian		Non-traditional	
Hours usually work in a week (Labor market)								
Female	36.298	(12.27)	34.585	(12.96)	37.717	(11.4)	40.155	(10.36)
Male	44.807	(10.49)	45.959	(10.35)	43.900	(10.11)	41.720	(13.87)
Income (in US dollars)								
Female hourly wage	6.380	(4.3)	6.027	(4.49)	6.619	(3.93)	7.727	(5.32)
Male hourly wage	7.510	(12.72)	7.208	(8.52)	7.699	(16.13)	8.825	(10.31)
Non-labor income/1000	0.064	(0.46)	0.082	(0.45)	0.037	(0.45)	0.129	(0.65)
Distribution factors								
Non-labor income difference/1,000 (Male-Female, US dollars)	-0.003	(0.26)	-0.006	(0.23)	-0.003	(0.29)	0.025	(0.33)
Married (1=Yes)	0.512	(0.5)	0.545	(0.5)	0.475	(0.5)	0.539	(0.5)
Age								
Female	40.303	(9.39)	41.215	(9.43)	39.410	(9.27)	39.674	(9.32)
Male	42.650	(9.76)	43.664	(9.69)	41.622	(9.7)	42.326	(10.14)
Race (1=White)								
Female	0.864	(0.34)	0.864	(0.34)	0.863	(0.34)	0.876	(0.33)
Male	0.873	(0.33)	0.887	(0.32)	0.863	(0.34)	0.829	(0.38)
Education (in years)								
Female	12.000	(3.79)	11.440	(3.78)	12.480	(3.68)	13.088	(4.01)
Male	10.762	(3.74)	10.268	(3.79)	11.185	(3.58)	11.720	(4.09)
Household characteristics								
N. children of both partners	0.822	(0.89)	0.821	(0.89)	0.821	(0.87)	0.855	(1.05)
N. children of female partner	0.133	(0.44)	0.133	(0.44)	0.138	(0.45)	0.083	(0.34)
N. children of male partner	0.023	(0.19)	0.021	(0.19)	0.026	(0.19)	0.021	(0.14)
House Owner (1=Yes)								
Female	0.083	(0.28)	0.080	(0.27)	0.084	(0.28)	0.114	(0.32)
Male	0.064	(0.24)	0.064	(0.25)	0.063	(0.24)	0.073	(0.26)
Number of rooms	3.571	(1.1)	3.599	(1.09)	3.536	(1.07)	3.627	(1.43)
Region								
Montevideo and Metropolitan area[/a]	0.605	(0.49)	0.529	(0.5)	0.671	(0.47)	0.756	(0.43)
North[/b]	0.145	(0.35)	0.200	(0.4)	0.090	(0.29)	0.124	(0.33)
South[/c]	0.250	(0.43)	0.271	(0.44)	0.239	(0.43)	0.119	(0.32)
Respondent (1=Female)	0.669	(0.47)	0.736	(0.44)	0.614	(0.49)	0.513	(0.5)
Observations	4,276		2,087		1,996		193	

Notes: The table reports the mean and standard deviations of the main variables of the empirical analysis. The sample is restricted to working couples were both members are between 23 and 64 years of age.

[/a] Montevideo, Canelones and San José.

[/b] Artigas, Paysnadú, Río Negro, Rivera, Salto, and Tacuarembó.

[/c] Cerro Largo, Colonia, Durazno, Flores, Florida, Lavalleja, Maldonado, Rocha, Soriano, and Treinta y Tres.

3.4 Empirical results

3.4.1 Main findings

The system of unrestricted labor supply on equations (3.7) and (3.8) is estimated using seemingly unrelated regression models (SURE) to allow for possible correlation on between the errors of the two equations.[14] The labor supply regressions for all couples, split according to the gender roles in the division of domestic work, are presented in Table 3.2.

The correlation between hours of market work and own-wage rate is negative in all household types. A one-percentage-point increase in the own-wage rate of females in traditional and egalitarian homes is associated with a reduction of approximately 5.7 and 5.2 hours of market work per month respectively. This correlation in non-traditional homes is not significant. Moreover, a one-percentage-point increase in the own-wage rate of males is associated with a reduction of 5.5 hours in traditional homes, while this magnitude is 5.1 in egalitarian and 6.4 in non-traditional households.[15] Regarding the cross-wage effects, the wage rate of females does not have any significant effect on the labor supply of males. Instead, the wage rate of men significantly increases the labor supply of women in traditional homes. The non-labor income is negatively correlated with the labor supply of males and females, but is only non-significant in the case of males in egalitarian homes and females in non-traditional households.

The non-labor income differences between males and females and married versus cohabiting partners have opposite correlations with the labor supply of women and men. An increase of $1000 in the non-labor income difference between males and females shifts the bargaining power of males, reducing the hours of market work for males and increasing it for women. This correlation is higher among traditional than egalitarian households, and it is not significant in non-traditional homes.[16] Therefore, the *income pooling* hypothesis of the unitary model is rejected in the sample of all couples, as well as in traditional and egalitarian households.

The dummy variable that accounts for the difference between married and cohabiting couples shows that being married empowers women in the full sample, reducing their labor

[14] Unfortunately, the ECH does not have information on parental education, religion, or other valid instruments to correct the endogeneity of wages and non-labor income. The second wave of the EFHU's survey has these variables, but it does not report the labor supply of the partner nor the wage rates.

[15] Rapoport et al. (2011) estimate similar labor supply equations that showed significant negative own-wage effects for females and males, whereas the estimation presented in Chiappori et al. (2002) indicates significant negative own-wage effects for men and significant positive effects for women.

[16] Oreffice (2011) found similar results over same-sex and heterosexual couples.

supply and increasing that of men. This correlation is significant in the full and egalitarian samples, only significant for men in non-traditional homes, and not significant in traditional homes. Thus, the insurance effect of marriage against divorce is present when members of the couple are less attached to a traditional division of domestic work. In traditional households, there is no significant difference between being married and cohabiting. This finding may reflect that in homes with traditional gender roles, women will be more specialized, regardless of the nature of their union.

Regarding the other variables included in the estimations, the results are similar to those found in the literature of collective models. The labor supply of males and females increases with an increase in own-age, but this variable is only significant among males. Own-education has a positive and significant correlation with labor supply, while an increase in partner's education decreases own-labor supply, with this finding significant for men in traditional households. The number of joint children the couple has is associated with fewer hours in the labor market for both partners in egalitarian homes, while as the number of children belonging to only one partner increases, the number of hours in the labor market of the other partner decreases. Being white or the house owner has no significant correlation on the labor supply of men or women, whereas the number of rooms in the house has a positive and significant effect on the labor supply of both partners. Concerning regions, the number of hours in the labor market is significantly lower in the north of the country than in the metropolitan area of the capital; however, there is no robust evidence of significant differences between the latter and the southern region. In traditional homes, when the respondent's survey is completed by the female they work fewer hours than when the respondent is the male partner, and the opposite is observed when the response is supplied by men.[17]

The proportionality test presented in equation (3.9) is tested using the results of the estimations presented in Table 3.2. The test imposes that the marginal effects of each distributional factor on labor supplies are equal. Thus, each variable affects the labor supply functions only through the sharing rule. The last two rows of Table 3.2 report the corresponding χ^2 test and $p - values$, where the null is that this equality holds. The joint hypothesis of equality of these ratios cannot be rejected in any household type considered in the analysis.[18] These results imply that collective rationality cannot be rejected for the

[17]The number of children of each member of the couple (from previous partners) and household ownership are often used as distribution factors. Considering that in this study the effect of these variables is not statistically significant, they are included as controls.

[18]The proportionality test is approximate because the married variable is defined as a dummy.

whole sample or any of the sub-samples of households used in this paper.[19]

The estimated results on non-labor income difference between partners and being married reject in traditional and egalitarian homes the restriction imposed by the unitary model, according to which, distribution factors do not have significant effects on household behavior, i.e., $f_4 = f_5 = m_4 = m_5 = 0$. The unitary model also requires that the data do not reject the Slutsky symmetry restriction, given by: $f_2 = f_3 = m_1 = m_3 = 0$. A Wald test of equality on the parameters associated with distribution factors to zero and a Wald test of Slutsky symmetry are both rejected for traditional and egalitarian households at 1% significance; however, for non-traditional homes, the former test is rejected at 10% and the latter at 5%. Therefore, there is no robust evidence of which model is more suitable for non-traditional homes.[20]

[19] I also estimated the model with caring preferences, and I cannot reject the collective rationality hypothesis.

[20] These results may be due to the small sample size of this type of household.

Table 3.2: Estimation of the unrestricted collective model of labor supply.

	All couples		Traditional		Egalitarian		Non-traditional	
	Female	Male	Female	Male	Female	Male	Female	Male
Log of wage of female	-5.243***	0.204	-5.774***	0.662	-5.271***	-0.158	-2.056	2.205
	(0.437)	(0.370)	(0.641)	(0.504)	(0.621)	(0.545)	(1.765)	(2.230)
Log of wage of male	0.572	-5.323***	1.774***	-5.545***	-0.108	-5.141***	-1.883	-6.439***
	(0.411)	(0.351)	(0.631)	(0.500)	(0.558)	(0.492)	(1.589)	(2.037)
Non-labor income/1000	-1.754***	-1.351***	-1.679***	-2.087***	-1.708***	-0.688	-2.127	-5.385***
	(0.421)	(0.356)	(0.641)	(0.502)	(0.616)	(0.540)	(1.456)	(1.845)
Non-labor income	1.949***	-1.432**	1.987*	-3.634***	1.766*	-0.191	4.058	0.925
difference/1,000	(0.708)	(0.603)	(1.194)	(0.944)	(0.930)	(0.820)	(2.528)	(3.248)
Married (1=Yes)	-1.132***	1.010***	-0.745	0.426	-1.485***	1.020**	0.296	5.404**
	(0.419)	(0.352)	(0.637)	(0.497)	(0.568)	(0.494)	(1.804)	(2.236)
Age	0.287*	0.582***	0.306	0.735***	0.374	0.452**	0.333	0.580
	(0.169)	(0.144)	(0.253)	(0.204)	(0.237)	(0.205)	(0.645)	(0.818)
Age squared	-0.004*	-0.007***	-0.003	-0.008***	-0.005	-0.005**	-0.004	-0.007
	(0.002)	(0.002)	(0.003)	(0.002)	(0.003)	(0.002)	(0.008)	(0.009)
Race (1=White)	0.249	0.067	-0.229	-0.468	0.870	0.627	-1.393	-2.443
	(0.532)	(0.466)	(0.810)	(0.685)	(0.719)	(0.635)	(2.333)	(2.531)
Education of female	0.708***	-0.122**	0.744***	-0.144*	0.582***	-0.022	0.665***	-0.438
	(0.068)	(0.058)	(0.104)	(0.082)	(0.092)	(0.081)	(0.251)	(0.319)
Education of male	-0.019	0.226***	-0.036	0.324***	-0.001	0.067	-0.254	1.019***
	(0.065)	(0.056)	(0.099)	(0.078)	(0.090)	(0.079)	(0.254)	(0.322)
N. children both	-1.046***	-0.186	-1.275***	0.262	-0.921***	-0.696**	0.043	0.136
	(0.238)	(0.199)	(0.358)	(0.278)	(0.335)	(0.288)	(0.802)	(1.009)
N. children female	-0.195	-0.307	-0.625	0.297	0.171	-1.110**	-0.020	0.520
	(0.444)	(0.373)	(0.672)	(0.523)	(0.602)	(0.522)	(2.221)	(2.829)
N. children male	-1.194	0.828	-0.019	1.247	-2.780**	1.027	-1.945	-2.483
	(0.967)	(0.825)	(1.416)	(1.122)	(1.342)	(1.181)	(5.357)	(6.742)
House Owner (1=Yes)	1.071	0.072	0.401	0.072	1.411	0.072	2.001	0.072
	(0.678)	(0.636)	(1.043)	(0.891)	(0.922)	(0.900)	(2.388)	(3.789)
Number of rooms	0.816***	1.123***	0.768***	0.911***	0.962***	1.305***	0.114	1.465*
	(0.188)	(0.160)	(0.284)	(0.225)	(0.264)	(0.232)	(0.672)	(0.876)
North	-3.353***	-2.107***	-2.818***	-2.804***	-2.416***	-1.901**	-1.340	-4.861*
	(0.548)	(0.466)	(0.749)	(0.591)	(0.897)	(0.787)	(2.356)	(2.937)
South	-1.318***	-0.104	-0.420	-0.565	-1.852***	0.142	0.687	-5.077*
	(0.448)	(0.380)	(0.676)	(0.531)	(0.606)	(0.534)	(2.282)	(3.005)
Female respondent	-1.414***	1.011***	-1.372**	0.970*	-0.480	0.499	-2.573*	-0.590
(1=Yes)	(0.394)	(0.336)	(0.632)	(0.498)	(0.524)	(0.463)	(1.495)	(2.005)
Constant	30.432***	36.413***	26.988***	34.439***	31.009***	39.068***	37.945***	27.130
	(3.413)	(3.037)	(5.243)	(4.417)	(4.646)	(4.210)	(13.019)	(17.176)
Proportionality test $\chi^2(1)$	0.06		0.30		0.74		0.03	
$p-value$	0.80		0.58		0.39		0.87	
Observations	4,276		2,087		1,996		193	

Notes: ***; **; * significant at 1%, 5% and 10%. Estimated coefficient and robust standard errors in parenthesis.

3.4.2 Results of the Collective model of Labor Supply

This subsection presents the results and implications of the collective model for the whole sample, as well as for traditional, egalitarian and non-traditional homes. Table 3.3 shows the estimates of the restricted collective model when the equality in equation (3.9) is imposed, while Table 3.4 tests whether distribution factors are relevant for the sharing rule. Tables 3.5 and 3.6 present the derivatives of the sharing rule and the wage and income elasticities for each household type respectively.

The results reported in Table 3.3 imposes the condition $m_4/f_4 = m_5/f_5$, which states that marginal effects of the income difference between the members of the couple should be equal to the ratio of marginal effects of being married on labor supplies. The estimation of the restricted model yields similar results to those found in Table 3.2. In particular, the distribution factors do not significantly change when imposing the collective rationality assumption.[21]

Table 3.3: Estimation of the restricted collective model of labor supply.

	All couples		Traditional		Egalitarian		Non-traditional	
	Female	Male	Female	Male	Female	Male	Female	Male
Log of wage of female	-5.241***	0.206	-5.775***	0.656	-5.273***	-0.161	-2.354	2.290
	(0.437)	(0.37)	(0.641)	(0.504)	(0.621)	(0.545)	(1.767)	(2.23)
Log of wage of male	0.571	-5.320***	1.770***	-5.542***	-0.100	-5.122***	-1.787	-6.46***
	(0.41)	(0.351)	(0.631)	(0.5)	(0.558)	(0.492)	(1.598)	(2.037)
Non-labor income/1000	-1.741***	-1.335***	-1.686***	-2.091***	-1.609***	-0.555	-1.173	-5.650***
	(0.418)	(0.353)	(0.64)	(0.502)	(0.605)	(0.518)	(1.321)	(1.836)
Non-labor income	1.861***	-1.514***	2.183*	-3.550***	1.373*	-0.721	0.409	1.951
difference/1,000	(0.603)	(0.504)	(1.161)	(0.935)	(0.802)	(0.516)	(0.913)	(3.176)
Married (1=Yes)	-1.186***	0.965***	-0.360	0.586	-1.623***	0.852*	1.090	5.204**
	(0.367)	(0.305)	(0.316)	(0.436)	(0.548)	(0.453)	(1.741)	(2.232)
Observations	4,276		2,087		1,996		193	

Notes: ***; **; * significant at 1%, 5% and 10%. Estimated coefficient and robust standard errors in parenthesis.

An additional condition must hold to recover the sharing rule up to an additive constant, i.e., at least one of the following conditions must hold: $f_4/f_3 \neq m_4/m_3$ or $f_5/f_3 \neq m_5/m_3$. This restriction implies that the ratio of marginal effects of each distribution factor and non-labor income should be different between partners. As distribution factors only affect the labor supply equations through the sharing rule, the signs should be opposite for females and males; whereas the marginal effect of non-labor income in the labor supply should be positive since leisure is a normal good. Hence, if the distribution factor "ℓ" increases

[21] Table 3.A.1 in the Appendix presents the complete table.

the bargaining power of females, the ratio f_{s_ℓ}/f_3 should be positive and m_{s_ℓ}/m_3 should be negative. Table 3.4 contains the estimated difference between these ratios, standard errors (in round brackets), and the value of the $\chi^2(2)$ Wald test estimates (in squared brackets). Results indicate that the sharing rule can be recovered up to an additive constant, for the whole sample, traditional, and egalitarian homes. However, the sharing rule is not identified for non-traditional households.

Table 3.4: Proportionality tests of the effect of non-labor income difference and being married.

	All couples	Traditional	Egalitarian	Non-traditional
$f_4/f_3 - m_4/m_3$	-2.171***	-2.925***	-1.312	-1.736
	(0.707)	(1.035)	(1.363)	(1.399)
	[9.42]	[7.98]	[0.93]	[1.54]
$f_5/f_3 - m_5/m_3$	1.393***	0.648	2.354*	0.864
	(0.427)	(0.47)	(1.429)	(0.99)
	[10.64]	[1.90]	[2.71]	[0.76]
Observations	4,276	2,087	1,996	193

Notes: ***; **; * significant at 1%, 5% and 10%. Estimated difference of coefficients and standard errors in parenthesis. Test of equality of these ratios between females and males (χ^2 values) in squared brackets.

Table 3.5 reports the results of estimated coefficients and marginal effects for all couples and households with different gender role attitudes.[22] According to equation (3.10), the marginal change in one variable reflects its impact on the non-labor income that goes to the female partner after the bargaining process has been completed. An increase in the log wage rate of the females member of the couple significantly increases her intra-household bargaining power in traditional and egalitarian homes, while a shift in the log wage rate of the male member significantly decreases her bargaining power only in traditional households. When these marginal changes are computed with respect to the wage rate and evaluated using sample means, the female in traditional homes significantly decreases her bargaining power, leaving her with a lower non-labor income.

The non-labor income impact on the sharing rule is significant in traditional and egalitarian households. A marginal increase of \$1 in non-labor income is estimated to give around \$0.52 to women in the full sample, almost \$0.57 in traditional households and \$0.60

[22]The estimation for the non-traditional homes is illustrative, but it is not valid.

in egalitarian. Therefore, the share of non-labor income allocated to women in traditional homes is slightly smaller than in egalitarian, but the difference is not significant. Browning et al. (2014) show that when domestic work is introduced into the collective model, and if the production function of the household admits complementarities between the number of hours of domestic work of both members, the result of higher bargaining power of women results in more housework for both members. Although the estimations in this study do not account for an accurate measure of leisure, as it does not introduce the number of hours of domestic work, the evidence seems to be in line with this theoretical result.

In the case of the non-labor income difference between males and females, there are significant results in the full sample and in traditional households (as expected from the results presented in Table 3.4). An increase of $1000 in favor of men will translate into a transfer of $550 to males in the full sample and $735 in traditional homes. Once again, there are no significant differences between traditional and egalitarian homes and intra-household bargaining appears to be highest for egalitarian women. Finally, being married has a positive and significant effect in the full sample and in egalitarian homes. This result implies that the females' share of non-labor income is higher among those who are married than those who cohabit.

Empirical evidence shows that bargaining power forces are relevant to intra-household bargaining in the whole sample, as well as in traditional and egalitarian homes. Regarding non-labor income, the magnitude is comparable to those found in the literature. For instance, Chiappori et al. (2002) found for a sample of households of United States, that a marginal increase of $1 increases the bargaining power of women by almost $0.68, and similar results are found in France by Rapoport et al. (2011). Oreffice (2011) using data from the US Census of the 2000s, estimates a marginal effect of non-labor income of $0.33 in heterosexual married couples and $0.56 in heterosexual cohabiting couples. Campaña et al. (2018) show that the marginal effect of non-labor income in the bargaining power of women is positive in Colombia, whereas it is negative in Mexico.

The results of the present study also suggest the rejection of the income pooling hypothesis, as has been found in many studies in for developed countries (Schultz, 1990; Lundberg et al., 1997; Lyssiotou, 2017; Oreffice, 2014). Oreffice (2011) uses a similar definition of the non-labor income difference between partners, and she rejects the income pooling hypothesis for heterosexual cohabiting and married couples for the United States. The overall magnitude of non-labor income difference in Uruguay is of a similar magnitude to that found for cohabiting couples in the United States.

Despite the diverging patterns in marriage and cohabitation, most papers that employ collective labor supply models do not distinguish between these types of unions, and those that do, analyze results on sub-samples of cohabiting and married couples (Oreffice, 2011). This paper uses marriage as a distribution factor to directly account for the differences in bargaining power among these types of unions. Blau and Kahn (2007) analyze the changes in labor supply of married women in the United States. They show that the higher propensity for cohabitation does not account for the fall in labor supply elasticities of married women.

Moreover, since marriage significantly increases the bargaining power of women in egalitarian households but not in traditional, self-selection into marriage does not appear to be a relevant issue.[23] Being married compensates for more specialization in egalitarian households by providing additional non-labor income to women after the bargaining process.

The age gap between males and females, the difference in educational attainment between partners, and sex ratios have all been explored as additional distribution factors. No significant effect was found of any of these variables on the labor supply of Uruguayan couples.

[23]Figure 3.A.1 in the Appendix shows that the main differences in participation rates and hours of market work are more associated with gender-role attitudes in the division of domestic work than with being married or cohabiting. The gap in labor participation rates among women with different social norms in the division of household chores reaches 20 percentage points, while the gap in the number of hours of market work is on average 8 hours per week. On the other hand, there are no significant differences in these variables between married and cohabiting women.

Table 3.5: Sharing rule coefficients and marginal effects for total, egalitarian and traditional couples.

	All couples		Traditional		Egalitarian		Non-traditional	
	Coeff.	$\partial\phi/\partial$Variable$^{/a}$	Coeff.	$\partial\phi/\partial$Variable$^{/a}$	Coeff.	$\partial\phi/\partial$Variable$^{/a}$	Coeff.	$\partial\phi/\partial$Variable$^{/a}$
Log of wage of female	0.075	0.010	1.943***	0.018	1.978*	-0.015	-212.7	5.6
	(0.136)	(0.018)	(0.711)	(0.015)	(1.082)	(0.051)	(41,552)	(1,088)
Log of wage of male	-0.169	-0.019	-0.596**	-0.067**	0.038	0.004	-161.0	-18.3
	(0.129)	(0.015)	(0.29)	(0.033)	(0.21)	(0.024)	(31,527)	(3,572)
Non-labor income /1,000$^{/b}$	0.515***		0.567***		0.604**		-105.9	
	(0.131)		(0.185)		(0.290)		(20,761)	
Non-labor income$^{/b}$ difference/1,000	-0.550***		-0.735**		-0.515		36.9	
	(0.177)		(0.300)		(0.370)		(7,174)	
Married$^{/b}$	0.351***		0.121		0.609*		98.5	
	(0.107)		(0.098)		(0.329)		(19,133)	
Observations	4,276		2,087		1,996		193	

Notes: ***; **; * significant at 1%, 5% and 10%. Robust standard errors in parenthesis.
$^{/a}$ The marginal effects are computed with respect to w_f and w_m, not with respect to $\log w_f$ and $\log w_m$, evaluated at the sample means.
$^{/b}$ The sharing rule is linear on non-labor income and distribution factors, which leads to the equality between estimated coefficients and marginal effects.

Table 3.6 presents the wage and non-labor income elasticities in the restricted general collective model.[24] At the sample mean, the women's and men's elasticities to own wages are negative and significant, except for women in non traditional homes. Women's own-wage elasticities vary from -0.167 in traditional households to -0.140 in egalitarian; while men's range from -0.155 to -0.117. In summary, the own-wage elasticities are larger for women than men in traditional and egalitarian households, but the opposite is true in non-traditional homes.

The cross-wage elasticity of females in traditional households is positive and significant, suggesting that these females may have incentives to move on to a more egalitarian division of labor. Non-labor income elasticity of both members of the couple is negative and of a similar magnitude to -0.003 across household types, but it is not significant for males in egalitarian households or females in non-traditional homes.

The results regarding labor supply elasticities are similar to those found for Colombia and Mexico (Campaña et al., 2018). Most of the empirical evidence for developed countries found that own-wage elasticities are negative for males but positive for females, and that cross-wage elasticities are negative (Chiappori et al., 2002; Oreffice, 2014; Lyssiotou, 2017; Goussé et al., 2017; Bloemen, 2019).

[24] Estimates of the unrestricted model are not different from those presented in this section. Full estimations are available upon request.

Table 3.6: Wage and non-labor income elasticities.

	All couples		Traditional		Egalitarian		Non-traditional	
	Female	Male	Female	Male	Female	Male	Female	Male
w_f	-0.144***	0.005	-0.167***	0.014	-0.140***	-0.004	-0.059	0.055
	(0.012)	(0.008)	(0.019)	(0.011)	(0.016)	(0.012)	(0.044)	(0.053)
w_m	0.016	-0.119***	0.051***	-0.121***	-0.003	-0.117***	-0.044	-0.155***
	(0.011)	(0.008)	(0.018)	(0.011)	(0.015)	(0.011)	(0.04)	(0.049)
y	-0.003***	-0.002***	-0.004***	-0.004***	-0.002***	-0.0005	-0.004	-0.017***
	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.0004)	(0.004)	(0.006)
Observations	4,276		2,087		1,996		193	

Notes: ***; **; * significant at 1%, 5% and 10%. Robust standard errors in parenthesis.

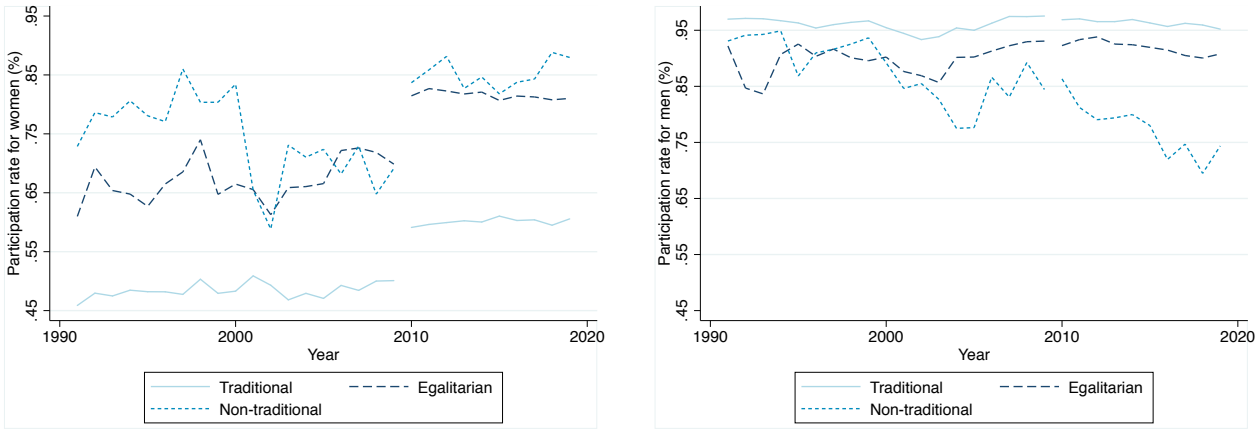
3.4.3 Selection into employment correction

The empirical analysis is extended to account for selection into employment, where one of the spouses specializes in household production. In the collective model, specialization takes place when the potential wage of one of the spouses is lower than his/her marginal productivity in housework and lower than his/her marginal rate of substitution between leisure and consumption. In this case, the distribution of domestic work between partners depends on the technology used in household production. Again, if there are complementarities to housework, a change in distribution factors that gives more bargaining power to women will increase the amount of time spent in household production for both (Browning et al., 2014).

The selection into employment correction allows the inclusion of households in which one of the two partners is out of the labor force. Households in which neither partner is in the labor force are not included. Figure 3.3a and 3.3b shows that the evolution of participation rate of females and males has narrowed among all household types. The participation rate of females has increased over the period, whereas that of males has remained stable in traditional and egalitarian households, and declined in non-traditional homes. As usual, the participation rate of females is lower than that of males. Differences in the participation rates between males and females are even more remarkable once the households' social norms regarding the division of domestic work are introduced.

The participation rate of women is the lowest, while men's is highest, in traditional homes. At the other extreme, the labor force participation of females in non-traditional households has increased, exceeding that of men at the end of the 2010s. Egalitarian homes

are in the middle of the distribution of participation rates for males and females.



(a) Women **(b)** Men

Figure 3.3: Labor force participation rates among women and men in Uruguay, 1991-2019.

Source: Author's calculations based on *Encuesta Continua de Hogares* from 1991 to 2019.
Note: Heterosexual couples between 23 and 64 years of age.

The Heckman correction is used to account for self-selection into employment and to compute fitted log wage. The wage equation is defined following a standard approach for human capital, which is common in collective models of labor supply (e.g.,Donni (2007), Oreffice (2011), and Donni and Matteazzi (2018)):

$$\log w_i = \mathbf{\Psi}' \mathbf{r^i} + u^i \tag{3.12}$$

where $i = f, m$, $\mathbf{r}$ includes each individual's education (measured in years), the square and cube of this variable, a second-order polynomial of the individual's age, a dummy for being white, dummies for the region of residence, a dummy variable if the individual have lived in the same administrative department of Uruguay since they were born, and a dummy if the household has a dishwasher. As noted by Donni (2007), the identification of the hourly wage requires that most of the variables included as controls in the market wage regression be excluded from the empirical model of the collective labor supply. The selection equation for each household member is given by:

$$d^i = \gamma_1 y + \gamma_2 s_1 + \gamma_3 s_2 + \gamma_4 child\_0to5 + \gamma_5 unempolyment + \mathbf{\Gamma}' \mathbf{x^i} + v^i \tag{3.13}$$

where d is an indicator variable that takes the value of one if the individual participates

in the labor market, and zero otherwise, y is non-labor income, s_1 is the difference in non-labor income between males and females, and s_2 is a dummy variable for married couples. The exclusion restrictions (i.e., variables that affect participation decision) are the number of children between 0 and 5 years in the household and the 2018 unemployment rate by department. The variables in $\mathbf{x}$ include dummies for age brackets and dummies for education levels. Selection into employment correction was estimated for all heterosexual couples in the sample between 23 and 64 years of age, and separately for traditional, egalitarian, and non-traditional couples.

The two-equation model presented in (3.12) and (3.13) is estimated by maximum likelihood, and the results are shown in Table 3.A.2 of the Appendix. The estimation results suggest that selection into employment is a relevant issue only among women, as the estimation of the correlation coefficient of the error terms of both equations is negative and statistically significant for women but not for men (ρ) in traditional and egalitarian homes, whereas it is significant and positive for women in non-traditional households. A negative (positive) correlation between the selection and wage equations implies that unobservable factors that increase the likelihood of employment, also decrease (increase) the wage rates.

Estimation of the system of labor supply equations (equations (3.7) and (3.8)) with selectivity-corrected log hourly wage for males and females is presented in Table 3.A.3 of the Appendix. The number of observations in the sample of couples with a traditional division of domestic work almost doubles, while in egalitarian and non-traditional homes it increased by approximately 50%. In traditional households, 87% of the increase in the sample is explained by women being out of the labor force, while this percentage declines to 66% in egalitarian and 32% in non-traditional homes. Therefore, the specialization pattern is more important among households with traditional gender roles. It should be noted that the distribution factors included in the participation equation have significant effects on the probability of employment, in the same direction as they do in the labor supply estimations. An increase in the non-labor income of men will reduce their likelihood of being employed and increase their partners', while being married decreases the probability of employment for women and increases that of men.

Table 3.A.3 presents the estimation results of the unrestricted collective model of labor supply. No significant differences are observed in the correlation between women's and men's labor supply and non-labor income, nor the distribution factors. However, there are changes in the signs and significance of own and cross-wage rates. Once selection into employment is taken into account, own-wage has a significant positive impact on men's

labor supply for those living in traditional and egalitarian households. Furthermore, the cross-wage effect is negative and significant for both household members in traditional homes and only for females in egalitarian and non-traditional homes. The model does not reject collective rationality or any proportionality test in any household type (Tables 3.A.3 and 3.A.4 of the Appendix). Therefore, the sharing rule is well defined, and it can be recovered up to an additive constant for all couples, traditional, egalitarian, and non-traditional households.

Table 3.7 presents the results of the estimated parameters and marginal effects of the sharing rule after correcting for selection. Given the differences in the estimation of the unrestricted model of labor supply after correcting wages for selection, it is reasonable to observe changes in the sharing rule in these variables. Due to the fact that there may be additional endogeneity problems, the following comments are related to the non-labor income and distribution factors.

The selection correction reduces the magnitude of an increase of $1 on non-labor income, but signs and significance remained the same. This result may be related to the fact that women who are out of the labor force may have less bargaining power than those who are working. As before, women living in households with traditional gender norms have less bargaining power. Differences in non-labor income between males and females are significant in the whole sample, and among household types. Again, women living in households more attached to traditional gender roles lose relatively more of their intra-household bargaining when this difference increases. Lastly, being married significantly increases the intra-household bargaining of women in egalitarian households and the overall magnitudes are similar to those obtained without the selection correction.

Table 3.7: Sharing rule coefficients and marginal effects for total, egalitarian and traditional couples with Heckman's MLE correction.

	All couples		Traditional		Egalitarian		Non-traditional	
	Coeff.	$\partial\phi/\partial$Variable$^{/a}$	Coeff.	$\partial\phi/\partial$Variable$^{/a}$	Coeff.	$\partial\phi/\partial$Variable$^{/a}$	Coeff.	$\partial\phi/\partial$Variable$^{/a}$
Log of wage	-0.605***	-0.109***	-0.643**	-0.121**	-0.161	-0.027	0.613	0.100
of female	(0.233)	(0.0417)	(0.292)	(0.055)	(0.339)	(0.0575)	(1.048)	(0.1716)
Log of wage	1.883***	0.303***	1.607**	0.266**	1.706***	0.265***	1.65*	0.257*
of male	(0.478)	(0.0768)	(0.797)	(0.1321)	(0.626)	(0.0973)	(0.908)	(0.1414)
Non-labor income	0.408***		0.306***		0.475***		0.528***	
/1,000$^{/b}$	(0.05)		(0.101)		(0.065)		(0.149)	
Non-labor income	-0.990***		-1.150***		-0.860***		-0.859***	
difference/1,000$^{/b}$	(0.087)		(0.176)		(0.106)		(0.235)	
Married$^{/b}$	0.144***		0.083		0.111*		0.226	
	(0.04)		(0.055)		(0.057)		(0.2)	
Observations	7,377		4,086		2,958		333	

Notes: ***; **; * significant at 1%, 5% and 10%. Robust standard errors in parenthesis.

$^{/a}$ The marginal effects are computed with respect to w_f and w_m, not with respect to $\log w_f$ and $\log w_m$.

$^{/b}$ The sharing rule is linear on non-labor income and distribution factors, which leads to the equality between estimated coefficients and marginal effects.

To sum up, the correction of the estimates with selection into employment changes magnitudes but no signs in the intra-household bargaining power of non-labor income and distribution factors. It remains true that the bargaining power of women is higher in those households less attached to traditional gender roles, even though the differences among household types are not statistically significant.

3.5 Conclusions

This paper has analyzed differences in the labor supply responses and intra-household bargaining of heterosexual couples in households with different gender role attitudes regarding the division of domestic work. Households were classified as traditional, egalitarian, and non-traditional using data from Uruguay and information on who mainly does household chores in couples where both members earn. More educated, younger, and richer individuals tend to have more egalitarian gender role attitudes, as they self-reported that both members do domestic work.

The collective model of labor supply with distribution factors was used to analyze the labor supply of couples. Results show that the labor supply of males and females is responsive to bargaining power shifts measured by differences in the non-labor income between partners and in the condition of being married, or not. Results indicate that a relatively richer individual (whether male of female) reduces their labor supply in traditional homes,

while in egalitarian households being married significantly reduces the labor supply of women and increases that of men. Additionally, in traditional and egalitarian households, the data does not suggest the rejection of the collective rationality hypothesis for any household type nor the proportionality constraints for at least one distribution factor. On the contrary, the income-pooling hypothesis is rejected by the data.

The derivatives of the sharing rule show that an increase in non-labor income translates to a higher transfer to the female partner regardless of the distribution of domestic work, and it is higher among those households less attached to traditional gender role attitudes. However, these differences are not statistically different from zero among household types. The bargaining power of women living in traditional homes is reduced when the non-labor income difference increases in favor of males, while being married increases her bargaining power in egalitarian households. The latter result suggests that insurance against divorce is a relevant booster of women's intra-household bargaining while reinforcing the specialization pattern. These results are robust to the selection into employment correction.

Social norms regarding the division of domestic work are not the only relevant variable to explain differences in intra-household bargaining, but it was shown that they can booster the distribution of power within households. Future research may combine time-use data and information on attitudes towards social norms regarding the distribution of domestic work to disentangle the influence of these variables in the collective labor market choices and intra-household bargaining.

Bibliography

Agarwal, B. 1997. "Bargaining"and Gender Relations: Within and Beyond the Household. *Feminist Economics*, 3(1):1–51.

Amábile, F., Bucheli, M., González, C., and Lara, C. 2021. Gender Differences in Domestic Work during COVID 19 in Uruguay. *FCS-Decon, Udlear*, (12/21).

Batthyány, K., Genta, N., Perrotta, V., Aguirre, R., and Ferrari, F. 2015. *Los tiempos del bienestar social: Género, trabajo no remunerado y cuidados en Uruguay*. MIDES-InMujeres, Udelar-FCS.

Bayot, D. and Voena, A. 2015. Prenuptial Contracts, Labor Supply and Household Investments. *Unpublished manuscript, Department of Economics, University of Chicago, Illinois*.

Bertrand, M., Cortes, P., Olivetti, C., and Pan, J. 2021. Social Norms, Labour Market Opportunities, and the Marriage Gap between Skilled and Unskilled Women. *The Review of Economic Studies*, 88(4):1936–1978.

Bertrand, M., Kamenica, E., and Pan, J. 2015. Gender Identity and Relative Income within Households. *The Quarterly Journal of Economics*, 130(2):571–614.

Blau, F. D. and Kahn, L. M. 2007. Changes in the Labor Supply Behavior of Married Women: 1980–2000. *Journal of Labor economics*, 25(3):393–438.

Bloemen, H. G. 2019. Collective Labor Supply, Taxes, and Intrahousehold Allocation: An Empirical Approach. *Journal of Business & Economic Statistics*, 37(3):471–483.

Blundell, R., Chiappori, P.-A., and Meghir, C. 2005. Collective Labor Supply with Children. *Journal of Political Economy*, 113(6):1277–1306.

Bourguignon, F., Browning, M., and Chiappori, P.-A. 2009. Efficient Intra-Household Allocations and Distribution Factors: Implications and Identification. *The Review of Economic Studies*, 76(2):503–528.

Browning, M., Bourguignon, F., Chiappori, P.-A., and Lechene, V. 1994. Income and Outcomes: A Structural Model of Intrahousehold Allocation. *Journal of Political Economy*, 102(6):1067–1096.

Browning, M., Chiappori, P.-A., and Weiss, Y. 2014. *Economics of the Family*. Cambridge University Press.

Cabella, W. and Fernández Soto, M. 2017. Evolución reciente de las uniones consensuales entre los jóvenes del uruguay. *Notas de Población*, Vol. 44, Nro 105. CEPAL.

Campaña, J., Giménez, J. I., and Molina, J. A. 2018. Efficient Labor Supply for Latin Families: Is the Intra-Household Bargaining Power Relevant?

Chiappori, P.-A. 1988. Rational Household Labor Supply. *Econometrica: Journal of the Econometric Society*, pages 63–90.

Chiappori, P.-A. 1992. Collective Labor Supply and Welfare. *Journal of Political Economy*, 100(3):437–467.

Chiappori, P.-A., Fortin, B., and Lacroix, G. 2002. Marriage Market, Divorce Legislation, and Household Labor Supply. *Journal of Political Economy*, 110(1):37–72.

Chiappori, P.-A. and Mazzocco, M. 2017. Static and Intertemporal Household Decisions. *Journal of Economic Literature*, 55(3):985–1045.

Donni, O. 2007. Collective Female Labour Supply: Theory and Application. *The Economic Journal*, 117(516):94–119.

Donni, O. and Matteazzi, E. 2018. Collective Decisions, Household Production, and Labor Force Participation. *Journal of Applied Econometrics*, 33(7):1064–1080.

Fernández, R. 2013. Cultural Change as Learning: The Evolution of Female Labor Force Participation over a Century. *American Economic Review*, 103(1):472–500.

Galván, E. 2021. Gender Identity and Quality of Employment. *Economica*.

Gemici, A. and Laufer, S. 2011. Marriage and Cohabitation. *New York University, mimeo*.

Goussé, M., Jacquemet, N., and Robin, J.-M. 2017. Marriage, Labor Supply, and Home Production. *Econometrica*, 85(6):1873–1919.

Greenwood, J., Guner, N., Kocharkov, G., and Santos, C. 2016. Technology and the Changing Family: A Unified Model of Marriage, Divorce, Educational Attainment, and Married Female Labor-Force Participation. *American Economic Journal: Macroeconomics*, 8(1):1–41.

Haddad, G. K. 2015. Gender Ratio, Divorce Rate, and Intra-Household Collective Decision Process: Evidence from Iranian Urban Households Labor Supply with Non-participation. *Empirical Economics*, 48(4):1365–1394.

Lundberg, S. 2010. The Changing Sexual Division of Labour. In *The Shape of the Division of Labour*. Edward Elgar Publishing.

Lundberg, S. and Pollak, R. A. 1993. Separate Spheres Bargaining and the Marriage Market. *Journal of Political Economy*, 101(6):988–1010.

Lundberg, S. J., Pollak, R. A., and Wales, T. J. 1997. Do Husbands and Wives Pool their Resources? Evidence from the United Kingdom Child Benefit. *Journal of Human Resources*, pages 463–480.

Lyssiotou, P. 2017. The Impact of Targeting Policy on Spouses' Demand for Public Goods, Labor Supplies and Sharing Rule. *Empirical Economics*, 53(2):853–878.

Oreffice, S. 2007. Did the Legalization of Abortion Increase Women's Household Bargaining Power? Evidence from Labor Supply. *Review of Economics of the Household*, 5(2):181–207.

Oreffice, S. 2011. Sexual orientation and household decision making.: Same-sex couples' balance of power and labor supply choices. *Labour Economics*, 18(2):145–158.

Oreffice, S. 2014. Culture and Household Decision Making. Balance of Power and Labor Supply Choices of US-born and Foreign-born Couples. *Journal of Labor Research*, 35(2):162–184.

Pearse, R. and Connell, R. 2016. Gender Norms and the Economy: Insights from Social Research. *Feminist Economics*, 22(1):30–53.

PNUD 2020. Informe sobre Desarrollo Humano 2020, Panorama General. *La próxima frontera. El desarrollo humano y el Antropoceno* `http://hdr.undp.org/sites/default/files/hdr_2020_overview_spanish.pdf`

Rapoport, B., Sofer, C., and Solaz, A. 2011. Household Production in a Collective Model: Some New Results. *Journal of Population Economics*, 24(1):23–45.

Schultz, T. P. 1990. Testing the Neoclassical Model of Family Labor Supply and Fertility. *Journal of Human resources*, pages 599–634.

Seiz, M. 2021. Equality in Confinement: Nonnormative Divisions of Labor in Spanish Dual-Earner Families during the COVID-19 Lockdown. *Feminist Economics*, 27(1-2):345–361.

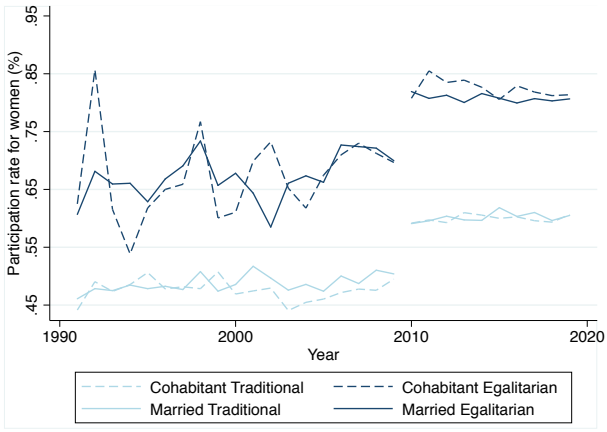
Vermeulen, F. 2005. And the Winner is... an Empirical Evaluation of Unitary and Collective Labour Supply Models. *Empirical Economics*, 30(3):711–734.

3.A Appendix for Chapter 3

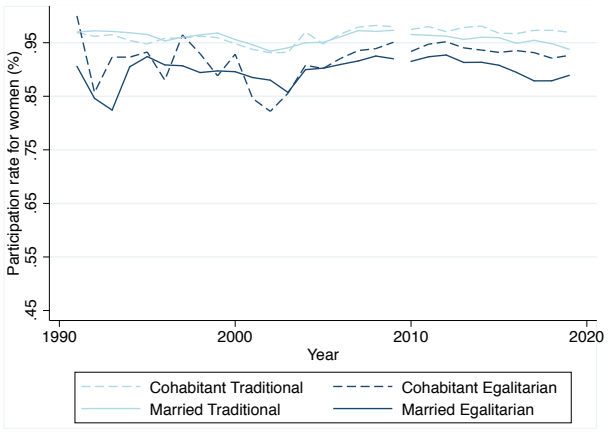
Table 3.A.1: Estimation of the restricted collective model of labor supply.

	All couples		Traditional		Egalitarian		Non-traditional	
	Female	Male	Female	Male	Female	Male	Female	Male
Log of wage of female	-5.241***	0.206	-5.775***	0.656	-5.273***	-0.161	-2.354	2.29
	(0.437)	(0.37)	(0.641)	(0.504)	(0.621)	(0.545)	(1.767)	(2.23)
Log of wage of male	0.571	-5.32***	1.77***	-5.542***	-0.1	-5.122***	-1.787	-6.46***
	(0.41)	(0.351)	(0.631)	(0.5)	(0.558)	(0.492)	(1.598)	(2.037)
Non-labor income/1000	-1.741***	-1.335***	-1.686***	-2.091***	-1.609***	-0.555	-1.173	-5.65***
	(0.418)	(0.353)	(0.64)	(0.502)	(0.605)	(0.518)	(1.321)	(1.836)
Non-labor income	1.861***	-1.514***	2.183*	-3.55***	1.373*	-0.721	0.409	1.951
difference/1,000	(0.603)	(0.504)	(1.161)	(0.935)	(0.802)	(0.516)	(0.913)	(3.176)
Married (1=Yes)	-1.186***	0.965***	-0.36	0.586	-1.623***	0.852*	1.09	5.204**
	(0.367)	(0.305)	(0.316)	(0.436)	(0.548)	(0.453)	(1.741)	(2.232)
Age	0.291*	0.584***	0.291	0.721***	0.373	0.468**	0.353	0.573
	(0.169)	(0.144)	(0.252)	(0.205)	(0.238)	(0.205)	(0.65)	(0.819)
Age squared	-0.004*	-0.007***	-0.003	-0.008***	-0.005	-0.006**	-0.005	-0.006
	(0.002)	(0.002)	(0.003)	(0.002)	(0.003)	(0.002)	(0.008)	(0.009)
Race (1=White)	0.309	0.082	-0.279	-0.389	1.082	0.628	-1.554	-2.471
	(0.534)	(0.467)	(0.815)	(0.688)	(0.722)	(0.637)	(2.349)	(2.533)
Education of female	0.709***	-0.123**	0.742***	-0.145*	0.583***	-0.023	0.687***	-0.444
	(0.068)	(0.058)	(0.104)	(0.082)	(0.092)	(0.081)	(0.252)	(0.319)
Education of male	-0.019	0.227***	-0.04	0.323***	0.00002	0.068	-0.269	1.022***
	(0.065)	(0.056)	(0.099)	(0.078)	(0.09)	(0.079)	(0.255)	(0.322)
N. children both	-1.043***	-0.187	-1.302***	0.255	-0.916***	-0.702**	-0.169	0.187
	(0.238)	(0.199)	(0.357)	(0.278)	(0.335)	(0.288)	(0.797)	(1.008)
N. children female	-0.205	-0.319	-0.552	0.329	0.135	-1.174**	-0.312	0.611
	(0.441)	(0.37)	(0.665)	(0.521)	(0.599)	(0.517)	(2.228)	(2.829)
N. children male	-1.2	0.831	0.008	1.284	-2.79**	1.025	-1.332	-2.676
	(0.967)	(0.824)	(1.415)	(1.122)	(1.342)	(1.182)	(5.371)	(6.742)
House Owner (1=Yes)	1.027	-0.065	0.404	-1.141	1.297	0.324	2.087	3.715
	(0.678)	(0.637)	(1.041)	(0.895)	(0.924)	(0.902)	(2.403)	(3.792)
Number of rooms	0.816***	1.124***	0.758***	0.905***	0.966***	1.307***	-0.018	1.508*
	(0.188)	(0.16)	(0.284)	(0.225)	(0.264)	(0.232)	(0.671)	(0.875)
North	-3.35***	-2.105***	-2.828***	-2.802***	-2.422***	-1.901**	-1.284	-4.889*
	(0.548)	(0.466)	(0.749)	(0.591)	(0.897)	(0.787)	(2.369)	(2.938)
South	-1.323***	-0.106	-0.408	-0.563	-1.865***	0.136	0.582	-5.029*
	(0.448)	(0.38)	(0.676)	(0.531)	(0.606)	(0.534)	(2.295)	(3.005)
Female respondent	-1.413***	1.003***	-1.374**	0.965*	-0.488	0.465	-2.811*	-0.526
(1=Yes)	(0.394)	(0.336)	(0.632)	(0.498)	(0.523)	(0.462)	(1.498)	(2.004)
Constant	30.295***	36.378***	27.425***	34.708***	30.798***	38.754***	38.373***	27.161
	(3.414)	(3.04)	(5.23)	(4.426)	(4.657)	(4.222)	(13.088)	(17.189)
Observations	4,276		2,087		1,996		193	

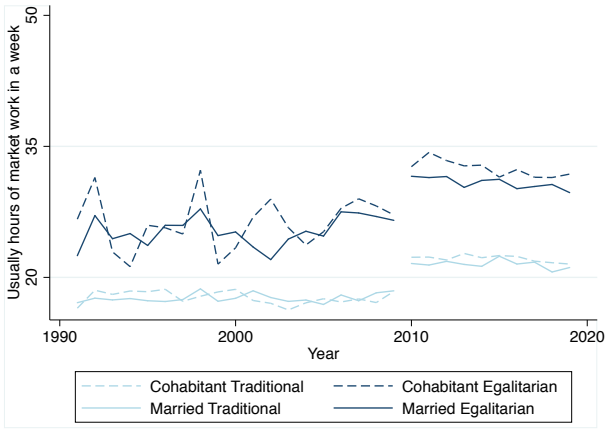
Notes: ***; **; * significant at 1%, 5% and 10%. Estimated coefficient and robust standard errors in parenthesis.



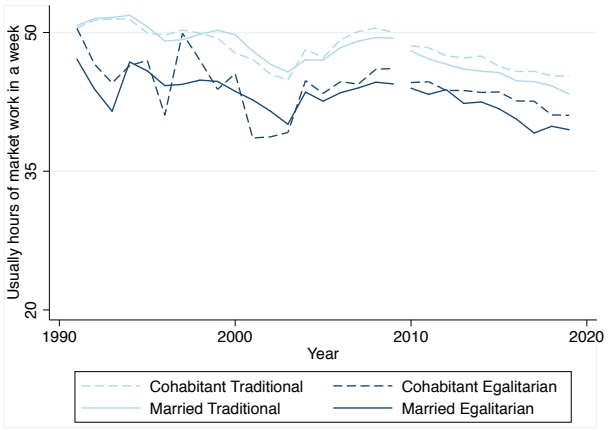
(a) Labor force participation of females



(b) Labor force participation of males



(c) Hours of market work female partner (not conditional on working)



(d) Hours of market work male partner (not conditional on working)

Figure 3.A.1: Hours of work in the labor market and participation rates in Uruguay, 1991-2019.

Source: Author's calculations based on *Encuesta Continua de Hogares* from 1991 to 2019. *Note:* Heterosexual couples between 23 and 64 years of age. The graphs do not include non-traditional households to avoid the variability of this group due to the small sample size.

Table 3.A.2: Estimation results for labor supply model with Heckman's MLE correction.

	All couples		Traditional		Egalitarian		Non-traditional	
	Female	Male	Female	Male	Female	Male	Female	Male
Wage equation								
Age	-0.000**	-0.000***	-0.000***	-0.000***	-0.000	-0.000**	-0.000	-0.001*
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Age squared	0.013**	0.022***	0.030***	0.028***	-0.001	0.015**	0.022	0.031
	(0.006)	(0.005)	(0.008)	(0.008)	(0.009)	(0.007)	(0.025)	(0.025)
Education	-0.192***	-0.093***	-0.163***	-0.121***	-0.242***	-0.067**	-0.138*	-0.248**
	(0.028)	(0.023)	(0.048)	(0.040)	(0.039)	(0.028)	(0.084)	(0.113)
Education squared	0.001***	0.001***	0.000	0.001**	0.001***	0.001***	0.001	0.002**
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.001)
Race (1=White)	0.020***	0.010***	0.019***	0.015***	0.022***	0.006***	0.017***	0.024***
	(0.002)	(0.002)	(0.004)	(0.003)	(0.003)	(0.002)	(0.006)	(0.008)
South	-0.001***	-0.000***	-0.001***	-0.000***	-0.001***	-0.000**	-0.000**	-0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
North	0.018	0.064***	0.030	0.051*	0.005	0.075***	0.067	0.092
	(0.018)	(0.017)	(0.026)	(0.027)	(0.027)	(0.023)	(0.081)	(0.081)
Constant	1.446***	0.989***	0.096	1.621***	0.096	1.356***	0.832	1.529***
	(0.189)	(0.150)	(0.194)	(0.241)	(0.194)	(0.249)	(0.740)	(0.581)
Participation equation								
Non-labor income/1000	-0.261***	-0.381***	-0.394***	-0.264***	-0.181***	-0.580***	-0.507**	-0.423***
	(0.035)	(0.049)	(0.061)	(0.072)	(0.045)	(0.082)	(0.204)	(0.156)
Non-labor income	0.965***	-1.016***	1.053***	-0.882***	0.852***	-1.329***	1.556***	-0.937***
difference/1,000	(0.055)	(0.062)	(0.094)	(0.087)	(0.071)	(0.110)	(0.372)	(0.176)
Married (1=Yes)	-0.167***	0.142***	-0.122*	0.032	-0.156***	0.252***	-0.331	0.137
	(0.036)	(0.053)	(0.064)	(0.079)	(0.046)	(0.087)	(0.239)	(0.196)
Age								
30-40	0.291***	0.019	0.301***	-0.003	0.301***	0.074	1.127***	0.228
	(0.052)	(0.103)	(0.085)	(0.140)	(0.070)	(0.187)	(0.355)	(0.347)
40-50	0.232***	-0.057	0.310***	0.002	0.249***	0.043	0.684*	-0.202
	(0.057)	(0.104)	(0.096)	(0.144)	(0.076)	(0.189)	(0.357)	(0.340)
50-60	0.167***	-0.242**	0.129	-0.266*	0.235***	-0.169	0.433	-0.178
	(0.065)	(0.108)	(0.107)	(0.147)	(0.085)	(0.195)	(0.398)	(0.378)
$\geq$60	-0.301***	-0.793***	-0.362**	-0.759***	-0.208*	-0.748***	-0.161	-0.877**
	(0.095)	(0.118)	(0.162)	(0.167)	(0.123)	(0.204)	(0.528)	(0.420)
Education								
High school not completed	0.357***	-0.015	0.270***	-0.123	0.379***	0.254***	0.100	0.079
	(0.045)	(0.058)	(0.085)	(0.099)	(0.055)	(0.087)	(0.326)	(0.222)
High school completed	0.885***	0.120	0.757***	0.138	0.855***	0.295**	0.565	0.214
	(0.056)	(0.079)	(0.099)	(0.125)	(0.071)	(0.126)	(0.366)	(0.275)
University not completed	0.939***	0.186	0.937***	0.204	0.836***	0.208	0.876*	0.921**
	(0.078)	(0.125)	(0.137)	(0.182)	(0.102)	(0.205)	(0.511)	(0.454)
University completed	1.598***	0.808***	1.453***	0.595***	1.548***	1.326***	1.272***	1.756***
	(0.063)	(0.115)	(0.110)	(0.154)	(0.082)	(0.231)	(0.410)	(0.475)
N. child 0 to 5 years old	-0.220***	0.064	-0.154***	0.152*	-0.251***	0.108	0.043	-0.350**
	(0.033)	(0.054)	(0.057)	(0.083)	(0.043)	(0.094)	(0.261)	(0.175)
Unemployment rate	-0.014**	0.009	0.010	-0.011	-0.012	0.024*	-0.042	-0.075*
	(0.007)	(0.010)	(0.015)	(0.018)	(0.008)	(0.015)	(0.050)	(0.041)
Constant	0.025	1.476***	0.096	1.621***	0.096	1.356***	0.832	1.529***
	(0.097)	(0.152)	(0.194)	(0.252)	(0.194)	(0.249)	(0.740)	(0.581)
ρ	-0.258***	-0.087	-0.299***	-0.161	-0.223***	-0.083	0.566**	0.399
	(0.059)	(0.057)	(0.083)	(0.076)	(0.091)	(0.089)	(0.288)	(0.271)
Observations	7,377		2,958		4,086		333	

Notes: ***; **; * significant at 1%, 5% and 10%. Estimated coefficients, standard errors in parenthesis. estimation.

Table 3.A.3: Estimation of the unrestricted collective model of labor supply, with Heckman's MLE correction.

	All couples		Traditional		Egalitarian		Non-traditional	
	Female	Male	Female	Male	Female	Male	Female	Male
Log of wage of female	2.370	-5.395**	-0.310	-5.159**	4.195	-1.338	11.620	5.344
	(2.898)	(2.108)	(3.366)	(2.264)	(4.714)	(3.469)	(8.850)	(10.161)
Log of wage of male	-9.660***	8.895***	-6.232**	5.178**	-9.995***	6.537*	-14.071**	5.153
	(2.248)	(2.286)	(2.865)	(2.575)	(3.493)	(3.845)	(6.200)	(8.491)
Non-labor income	-2.070***	-5.481***	-1.168**	-5.710***	-2.805***	-5.354***	-4.607***	-4.776***
/1000	(0.394)	(0.310)	(0.553)	(0.386)	(0.595)	(0.511)	(1.333)	(1.675)
Non-labor income	5.009***	-9.252***	4.511***	-9.493***	5.041***	-8.903***	7.483***	-8.120***
difference/1,000	(0.496)	(0.393)	(0.718)	(0.508)	(0.729)	(0.629)	(1.577)	(1.992)
Married (1=Yes)	-2.228***	0.935**	-1.956***	0.364	-2.112***	0.631	-0.720	3.952
	(0.484)	(0.378)	(0.660)	(0.459)	(0.723)	(0.616)	(2.096)	(2.547)
Age	2.005***	1.186***	2.022***	1.381***	1.951***	1.052***	1.400*	1.218
	(0.178)	(0.151)	(0.234)	(0.181)	(0.289)	(0.251)	(0.739)	(1.003)
Age squared	-0.024***	-0.016***	-0.024***	-0.018***	-0.024***	-0.015***	-0.019**	-0.018
	(0.002)	(0.002)	(0.003)	(0.002)	(0.003)	(0.003)	(0.008)	(0.011)
Race (1=White)	0.800	-0.896*	0.887	-1.296**	0.823	-0.381	-2.200	-3.453
	(0.594)	(0.507)	(0.801)	(0.638)	(0.914)	(0.802)	(2.616)	(2.935)
Education of female	1.345***	0.103	1.620***	0.143	0.847**	-0.094	0.403	-0.616
	(0.232)	(0.169)	(0.276)	(0.186)	(0.365)	(0.271)	(0.808)	(0.942)
Education of male	0.616***	-0.445**	0.316	-0.134	0.696**	-0.434	0.698	0.764
	(0.179)	(0.179)	(0.231)	(0.203)	(0.275)	(0.300)	(0.620)	(0.840)
N. children both	-2.452***	-0.399**	-3.014***	-0.166	-1.733***	-0.433	0.389	-2.173*
	(0.259)	(0.200)	(0.351)	(0.240)	(0.400)	(0.337)	(0.936)	(1.166)
N. children female	-1.202**	-0.862**	-1.728***	-0.686	-0.277	-1.402**	0.095	-0.131
	(0.470)	(0.367)	(0.613)	(0.427)	(0.740)	(0.626)	(2.435)	(3.098)
N. children male	-0.153	-0.329	-0.040	0.682	-1.443	0.178	3.108	-12.551
	(1.172)	(0.927)	(1.662)	(1.172)	(1.645)	(1.413)	(6.048)	(7.677)
House Owner (1=Yes)	1.464*	-0.153	0.699	-0.747	1.484	0.130	4.171	3.900
	(0.772)	(0.642)	(1.079)	(0.756)	(1.159)	(1.100)	(2.593)	(4.411)
Number of rooms	0.406*	0.821***	0.319	0.686***	0.706**	0.881***	-0.192	0.601
	(0.210)	(0.166)	(0.285)	(0.200)	(0.324)	(0.277)	(0.775)	(0.992)
North	-5.307***	-0.691	-4.033***	-2.005***	-2.872**	-0.212	2.060	-5.122
	(0.719)	(0.572)	(0.857)	(0.618)	(1.370)	(1.197)	(3.182)	(3.959)
South	-1.871***	0.373	-0.519	-0.876	-2.611***	1.059	5.510	-1.815
	(0.588)	(0.460)	(0.769)	(0.546)	(0.871)	(0.733)	(3.401)	(4.210)
Female respondent	-6.297***	4.117***	-5.977***	2.528***	-4.219***	3.528***	-3.924**	6.627***
(1=Yes)	(0.464)	(0.366)	(0.691)	(0.484)	(0.659)	(0.568)	(1.730)	(2.248)
Constant	-14.458***	15.344***	-19.682***	17.006***	-10.505**	16.584***	7.606	-4.804
	(3.533)	(3.069)	(5.008)	(3.885)	(5.337)	(4.827)	(14.623)	(20.695)
Proportionality test $\chi^2(2)$	3.66		0.72		0.52		0.08	
$p-value$	0.06		0.39		0.47		0.78	
Observations	7,377		2,958		4,086		333	

Notes: ***; **; * significant at 1%, 5% and 10%. Estimated difference of coefficient and standard errors in parenthesis. Regressions are corrected for sample selection.

Table 3.A.4: Proportionality tests of the effect of non-labor income difference and being married, with Heckman's MLE correction.

	All couples	Traditional	Egalitarian	Non-traditional
$f_5/f_3 - m_5/m_3$	-4.108	-5.526	-3.460	-3.324
	(0.488)	(1.864)	(0.458)	(0.946)
	[70.95]	[8.78]	[57.12]	[12.35]
$f_6/f_3 - m_6/m_3$	1.247	1.739	0.871	0.984
	(0.328)	(0.993)	(0.337)	(0.804)
	[14.47]	[3.06]	[6.66]	[1.50]
Observations	7,377	2,958	4,086	333

Notes: ***; **; * significant at 1%, 5% and 10%. Robust standard errors in parenthesis. Test of equality of these ratios between females and males (χ^2 values) in squared brackets.